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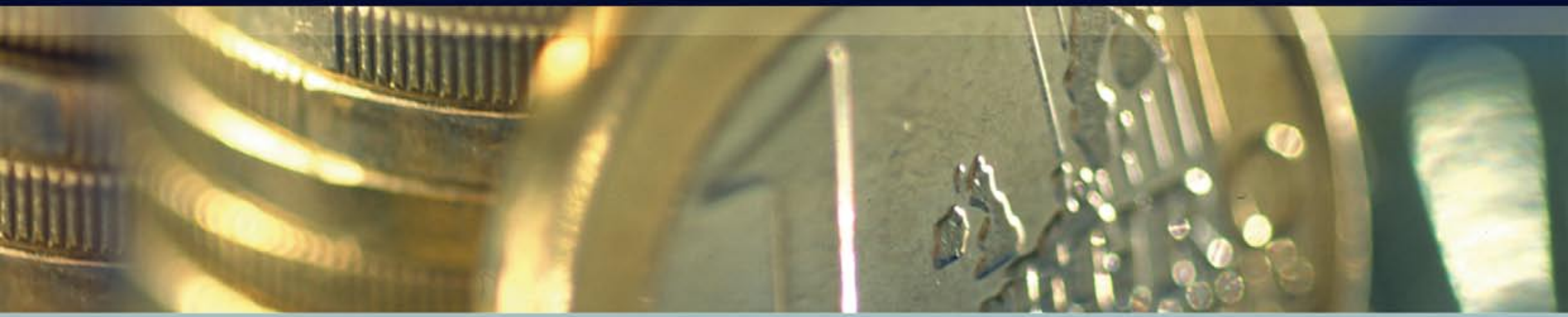

Time-varying integration, the euro and international diversification strategies

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Time-Varying Integration, the euro and International Diversification Strategies [1]

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Time-Varying Integration, the euro and International Diversification Strategies

Abstract

This paper investigates the impact of globalization and integration on the relative benefits of country and industry diversification. Unlike previous models, our factor model allows asset exposures to vary with both structural changes and temporary fluctuations in the economic and financial environment. First, we find that globalization and integration have lead to a gradual convergence of country to industry betas, especially in Europe. Second, not accounting for time-varying factor exposures leads to substantial biases in measures of country and industry risk. Third, even though the edge has structurally decreased, geographical diversification continues being superior to industry diversification.

JEL Classification: G11, G12, G15, C32, F37

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Time-Varying Integration, the euro and International Diversification Strategies

Executive Summary

The last decades have witnessed a rather dramatic increase in both real and financial globalization. Progress has been particularly impressive in Western Europe, a region that has gone through a unique period of economic, financial, and monetary integration. In 1999, this process culminated in the introduction of a single currency, the euro, in 11 countries (four more have joined since then).

The aim of this paper is to investigate whether globalization, regional integration, and the introduction of the euro have made local European (and non-European) equity markets more exposed to global and regional shocks. This is interesting for a number of reasons. First, previous research has shown that there is a strong positive link between common market exposures and financial integration, i.e. we can derive both the level and dynamics of market integration from these common factor exposures. Second, the results in this paper may yield interesting insights in the changing nature of shock transmission in European (global) equity markets, and the possibilities for risk sharing. In this paper, we will predominantly focus on the latter.

Our empirical model relates local equity market returns to both global and regional equity market shocks. This allows us to distinguish between global and regional integration. Our paper differs from previous work in the way time variation in market exposures, or betas, is modelled. First, we make the betas a linear function of three structural instruments, namely trade integration, industry structure alignment, and a euro dummy. Trade integration, measured as the ratio of imports plus exports over GDP, is not only a measure of economic integration, it is also strongly positively correlated with financial integration. Industry alignment measures to what extent the industrial structure of a country has become more similar to that of the global (regional) market. Last but not least, we allow the exposures to change after the introduction of the euro, a particularly important structural break. As we argue in the paper, we expect a positive relationship between global (regional) factor exposures and all three structural instruments. Second, we allow the beta intercept to switch between a high and a low value according to a latent

regime variable. This regime-switching intercept can either introduce cyclicality in the market betas, or proxy for structural changes in market betas not accommodated for by our three structural instruments.

We estimate our model on a set of 21 developed equity markets, of which 14 European. We observe a number of interesting patterns. First, we find that Europe as a region has become substantially more exposed to global market shocks, suggesting that Europe as a whole has become better integrated with world equity markets. Second, the exposures of local European equity market shocks to both global and regional shocks have increased structurally over time. The increase is larger for the global than for the regional market beta (respectively with 38 and 28 percent), suggesting that global integration is at least as important as regional integration. Surprisingly, we observe only minor differences between euro area and other European markets. While we also observe structural increases in market betas for non-European countries, they tend to be smaller in magnitude.

We are particularly interested in whether the euro has had an effect on market exposures. While we do not find a euro effect for Europe as a region, at the country level, we find a significant euro effect on the global market betas of 12 of the 21 countries. Except for Austria and Finland, the euro has resulted in an upward shift in global market betas that is not only statistically but also economically relevant (on average, + 0.14). Surprisingly, the effect is on average larger for countries not part of the euro area than for those that are (0.23 versus 0.13). Equally surprising is that there is hardly any effect of the euro on the regional market betas (on average, -0.02).

Next, we investigate to what extent the structurally-driven increase in common market exposures is associated with a structurally-driven decrease in asset-specific risk. For Europe as a whole, we find a decrease in region-specific volatility of about 3.4 percent, which corresponds to about 35 percent as a proportion of total volatility. At the country level, we observe a structural decrease in nearly all countries, with on average 2.85 percent in absolute value, or 20 percent as a proportion of total volatility. Interestingly, country-specific volatility is significantly lower after the introduction of the euro, especially in Europe.

The gradual increase in market betas and the corresponding decrease in country-specific risk have resulted in a structural increase in average cross-market correlations. For our entire sample of 21 countries, correlations have increased from about 40 percent in the 1970s to about 65 percent in the 2000s. The increase is substantially larger in continental Europe, to about 80 percent.

The positive conclusion from our paper is that equity market integration in Europe has increased substantially over time, not only within Europe, but also with other markets. The euro has strengthened this integration process. The negative conclusion is that while investors can now freely invest all over Europe (World), integration has structurally increased cross-market correlations, hereby reducing possibilities for risk sharing, at least between developed equity markets.

I Introduction

The last decades have witnessed a rather dramatic increase in both real and financial globalization. At a regional level, integration was often strengthened through the formation of free-trade areas or currency unions, with the introduction of the euro as its most visible example. Investors, especially in developed markets, now hardly face any direct or indirect barriers anymore to international investment, and should now be fully able to reap the large benefits of international diversification (see Solnik (1974)). Unfortunately, a large literature[2] has documented that equity market correlations tend to rise considerably exactly when markets become increasingly economically and financially integrated, hereby reducing the potential benefits of international diversification. The aim of this paper is to quantify the time-varying benefits of international diversification as more and more markets become increasingly integrated with world equity markets. We are particularly interested in whether integration has put an end to the traditional dominance of geographical over industry diversification potential (see e.g. Heston and Rouwenhorst (1994) and Griffin and Karolyi (1998)), as was recently suggested by Baca, Garbe, and Weiss (2000), Cavaglia, Brightman, and Aked (2000), Ferreira and Gama (2005), Campa and Fernandes (2006), and Eiling, Gerard, and de Roon (2007).

The extent to which the shift from geographical to industry diversification is permanent or not will ultimately depend on whether its underlying causes are structural rather than temporary. While previous papers have suggested that globalization and market integration are likely candidates to explain such a permanent shift, no paper has yet explicitly quantified their effects on diversification benefits. To fill this void in the literature, we estimate a dynamic factor model on index returns from 4 regions, 21 developed countries, and 18 global industries over the period 1973-2007. Based on the estimates of this factor model, we calculate and compare two popular indicators of diversification potential. First, we use our factor model to decompose total country and industry risk into a component that is due to common factor exposures, and a component that can be diversified respectively across countries and industries. Second, we calculate average model-implied cross-country and industry correlations over time. If globalization and market integration have indeed reduced international diversification benefits, one would expect a structural

[2] See Longin and Solnik (1995), Bekaert and Harvey (2000), Carrieri, Errunza, and Sarkissian (2004b), Goetzmann, Rouwenhorst, and Li (2005), and Bekaert, Hodrick, and Zhang (2005b) amongst many others.

decrease in average country-specific risk, and an increase in cross-country correlations. Similarly, for the shift from country to industry diversification to be permanent or not, one should observe a structurally-driven increase in industry-specific relative to country-specific volatility, and a convergence of average cross-country correlations to cross-industry correlations.

The main difference between our and previous approaches is that we explicitly allow both factor exposures and asset-specific volatilities to vary over time. Previous work has typically imposed strong restrictions of constant and unit factor exposures. The dominant empirical model in this literature, the Heston and Rouwenhorst (1994) model, assumes unit exposures with respect to a market and a large number of country and industry factors[3]. Similarly, in their extention of the Campbell, Lettau, Malkiel, and Xu (2001) model, Ferreira and Gama (2005) investigate the evolution of global, country, and local industry risk over time using a factor model that implicitly assumes country and industry returns to have a unit exposures to global market shocks. However, as we show further in this paper, not accounting for the fact that market betas are different from one and time-varying may lead to serious biases in measures of country and industry diversification potential.

We see a number of reasons why market betas may indeed differ from one and vary through time. First, there is a large literature documenting that global factor exposures increase from values close to zero to values closer to one as markets become better and better integrated (see e.g. Bekaert and Harvey (1997), Bekaert and Harvey (2000), Ng (2000), Fratzscher (2002), Baele (2005)). Second, industry betas may change through time as to reflect changing industry characteristics and/or regulation. For instance, market betas of many European (US) banks increased after the Second Banking Directive in 1989 (Gramm-Leach-Bliley Act in 1999) allowed banks to also enact in non-traditional banking activities, such as insurance and investment banking (see e.g. Baele, De Jonghe, and Vennet (2007a)). Third, even in the absense of structural change, both industry and country betas may still vary substantially over the business cycle[4].

[3]Brooks and Del Negro (2006) extend the Heston and Rouwenhorst to allow asset-specific factor exposures. However, their approach does not allow for time-varying factor exposures.

[4]For instance, evidence in Ferson (1989), Ferson and Harvey (1991), Ferson and Harvey (1993), Ferson and Korajczyk (1995), Jagannathan and Wang (1996), and more recently Santos and Veronesi (2004) indicates that betas are a function of economic state variables. While Ghysels and Jacquier (2005) find

To decompose global industry and regional risk into a systematic and diversifiable component, we include global market returns as an obvious first factor. At the country level, we include both a global and a regional factor. Bekaert et al. (2005b) show that such a model adequately describes comovement between country returns. An additional advantage of this two factor approach is that we can measure how much of country-specific risk can be eliminated by diversifying within the region. As argued before, we mainly differ from previous approaches in the way we model factor exposures and conditional volatilities. We model the market betas as a linear function of two structural instruments - proxying for time-varying market integration and development - and a latent regime variable. The regime-switching intercept could either capture cyclical variation in market betas or structural changes not accommodated by our structural instruments. Second, we also add structural economic variables to the traditional Asymmetric GARCH specification for the conditional asset-specific volatility[5]. By not including this channel, previous volatility spillover papers inherently assume that structural increases in market betas lead to a systematic increase in the assets' total risk, an implication that - at least at the country and industry level - contradicts previous evidence (see e.g. Ferreira and Gama (2005)). By including structural instruments both in the beta and asset-specific volatility specifications, we allow structural changes in the betas to have a compensating effect on asset-specific volatility. Finally, we also test whether the introduction of the euro - a particularly relevant structual break - has had an additional effect on both common market shocks and asset-specific volatilities.

Our model estimates reveal a number of interesting patterns. First, the market exposures of both countries and industries tend to deviate substantially and for long periods from one and/or from its unconditional value. As a consequence, models assuming unit/constant factor exposures risk introducing large biases in standard diversification measures. Second, the dynamics of market betas differs substantially between countries and industries. For most countries, especially in Europe, we find a structurally-driven increase in market betas towards one. Instead, global industry betas mostly follow a (counter)cyclical pattern, and are relatively unaffected by structural changes. The structural increase in market betas

only a limited role for macro-economic or firm-specific variables, they do confirm that substantial time variation in market betas.

[5]Our modeling approach essentially decomposes conditional betas and volatilities into a slowly-moving structurally-driven component and a (higher-frequency) transitionary component (see e.g. Engle and Rangel (2005) and Engle, Ghysels, and Sohn (2006) for similar approaches).

for countries is not only statistically but also economically relevant, and amounts often to more than 50 percent. Third, we find that the introduction of the euro has had a surprisingly large effect on market exposures, even outside Europe. Fourth, we show that structural increases in market betas tend to be associated with structural decreases in the level of country-specfic risk. As for market betas, this effect is not only statistically but also economically relevant, and amounts to on average 20 percent of country-specific volatility.

We then investigate the implications of our model estimates for optimal diversification strategies. First, we show that the time variation in factor exposures is sufficiently large to create substantial biases in measures of average country and industry-specific risk that are based on models assuming constant or unit exposures. In our sample, we find biases in average country and industry-specific risk of more than 25 percent. The bias in industry-specific risk is generally below 10 percent, but rises to nearly 30 percent in the period corresponding to the TMT bubble. In particular, we show that unit and constant beta studies have overstated the benefits of geographical diversification especially in the early 1970s and between 1985-1995, while overstating the benefits on industry diversification at the end of the 1990s. Consequently, our results indicate that the strong rise in the relative potential of industry diversification at the end of the 1990s may not have been as large and clear-cut as recently suggested by a number of papers. Second, after correcting for these biases, we find that, over the last 30 years, average country-specific risk was typically higher than average industry-specific risk. When time-varying betas are accounted for, the edge of geographical over industry diversification is, however, substantially lower than in unit/constant beta studies. Fourth, we find that globalization and further regional integration have lead to a gradual but strong increase in cross-country correlations. By the end of the 1990s, average cross-country and industry correlations were roughly at the same level. Fifth, similar to previous work, we do find a substantial increase (decrease) in average industry-specific volatility (average cross-industry correlations) at the start 2000s. Contrary to Brooks and Del Negro (2004), we show that this rise is partly but not fully an artifact of the buildup and burst of the TMT bubble at the end of the 1990. Last but not least, we find that as from 2003 on, cross-industry correlations again rose above cross-country correlations, while average country-specific volatility again rose above average cross-industry volatility. We conclude that while the benefits of geographical diversification have gradually decreased with globalization and integration, they are still substantial, to the extent that geographical diversification yields still larger risk reduction benefits than industry diversification.

The remainder of this paper is organized as follows. Section II develops a structural regime-switching methodology that allows both betas and idiosyncratic volatility to vary with both temporary and structural factors. Section III describes the stock return data as well as the structural instruments used in the estimation process. Section IV discusses the estimation results from the volatility spillover model. In Section V, we analyze the portfolio implications of our model estimates. Finally, Section VI provides conclusions.

II Structural Regime-Switching Volatility Spillover Model

In this section, we develop a volatility spillover model that decomposes total volatility at the regional, country, and global industry level in a systematic and an idiosyncratic component. To correctly separate systematic and idiosyncratic risk, we allow the exposures to global and regional market shocks to vary with both structural changes and temporary fluctuations in the economic environment.

A Model Specification

Let $r_{gi,t}$ represent the excess return of global industry gi, which we decompose as follows:

$$r_{gi,t} = \mu_{gi,t-1} + \beta^{w}_{gi,t-1}\varepsilon_{w,t} + \varepsilon_{gi,t} \tag{1}$$

where $\mu_{gi,t-1}$ represents the global industry's expected return, $\varepsilon_{w,t}$ the global equity market shock, and $\varepsilon_{gi,t}$ the global industry-specific shock. The global market shocks are calculated as $\varepsilon_{w,t} = r_{w,t} - \mu_{w,t-1}$, where $r_{w,t}$ is the excess market return and $\mu_{w,t-1}$ its conditional mean. The conditional dependence of global industry shocks on global market shocks is determined by $\beta^{w}_{gi,t-1}$. Similarly, the excess return on a regional market is specified as follows:

$$r_{reg,t} = \mu_{reg,t-1} + \beta^{w}_{reg,t-1}\varepsilon_{w,t} + \varepsilon_{reg,t} \tag{2}$$

where $\mu_{reg,t-1}$ constitutes region reg's expected return, $\beta^{w}_{reg,t-1}$ the region's beta with respect to global market shocks, and $\varepsilon_{reg,t}$ the equity market shock specific to region reg. Both the specification for global industry and regional market shocks correspond to the one-factor volatility spillover model of Bekaert and Harvey (1997). Finally, the excess return on the market index of a particular country c is given by:

$$r_{c,t} = \mu_{c,t-1} + \beta^{w}_{c,t-1}\varepsilon_{w,t} + \beta^{reg}_{c,t-1}\varepsilon_{reg,t} + \varepsilon_{c,t} \tag{3}$$

where $\mu_{c,t-1}$ represents country c's expected return, $\varepsilon_{w,t}$ and $\varepsilon_{reg,t}$ respectively the global and regional equity market shocks, and $\varepsilon_{c,t}$ the country-specific shock. The conditional exposure of local market shocks to global and regional market shocks is governed by respectively $\beta^{w}_{c,t-1}$ and $\beta^{reg}_{c,t-1}$. This model corresponds to the two-factor volatility spillover model of Ng (2000), Bekaert, Harvey, and Ng (2005a), and Baele (2005). The main advantage of this specification is that it allows to differentiate between global and regional integration.

The existing spillover literature has made the global (regional) market betas time-varying by conditioning them on some structural information variables (see e.g. Bekaert and Harvey (1997), Ng (2000) and Bekaert et al. (2005a)) or on a latent regime variable (see Baele (2005)). Both specifications on their own, however, risk missing important features of the beta dynamics. While the first approach allows betas to change with structural changes in the economic and financial environment, it cannot accommodate either cyclical or purely temporal variation in the betas. The second approach does allow betas to fluctuate over time, but is less suited to deal with *permanent* changes in market betas.

A first methodological innovation of this paper is that we condition the global (regional) market betas both on a number of structural economic instruments *and* on a latent regime variable, hereby allowing for both structural changes and temporal fluctuations in market betas[6]. The general specification for the global (regional) market betas is given by:

$$\beta^{om}_{rm,t} = \beta^{om}_{rm}(S_{rm,t}) + \beta^{om}_{rm}X_{rm,t-1} \tag{4}$$

where subscript $om = \{w, reg\}$ indicates the market (industry) where the shocks originate, and subscript $rm = \{reg, c, gi\}$ the receiving market (industry). The latent regime

[6]Note that the above specification implies a number of cross-sectional restrictions on the market betas. For instance, a value-weighted average of global market industry, region, and country betas should equal one at each point in time. While our estimation procedure (see Section II.B) does not permit us to impose these restrictions, we will test ex post whether our estimates violate these restrictions or not (and we find they do not violate them, see Section IV.C).

variables $S_{rm,t}$, the structural instruments $X_{rm,t-1}$, and the sensitivity to the instruments are all allowed to be different for each receiving market or industry. To limit the parameter space, we impose that the dynamics of the global and regional market beta intercept is driven by the same (but country-specific) latent regime variable. This does not mean, however, that we impose the global and regional betas to have the same evolution over time, as both betas are still allowed to have a specific exposure to global (region)-specific structural instruments.

A second methodological contribution of this paper is that we allow structural changes in the market betas to have a feedback effect on the level of asset-specific volatility. For instance, for a constant level of total volatility, one would expect an increase in market betas to have a dampening effect on the level of idiosyncratic volatility. Methodologically, we assume the asset-specific shocks to be distributed as follows:

$$\varepsilon_t \sim N\left(0, \Xi_t\right)$$

where $\Xi_t = diag\left(h_{z,t}\right)$ and $z = \{w, reg, c, gi\}$. We assume hence that all covariance between the asset returns is accommodated through the respective (time-varying) betas and factor volatilities[7]. In its most general form, the conditional volatility for asset z is given by:

$$h_{z,t} = \varkappa Q_{z,t-1} + \psi_{z,0}\left(S_{z,t}^V\right) + \psi_{z,1}\left(S_{z,t}^V\right)\varepsilon_{z,t-1}^2 + \psi_{z,2}\left(S_{z,t}^V\right)h_{z,t-1} + \psi_{z,3}\left(S_{z,t}^V\right)\varepsilon_{z,t-1}^2 I\{\varepsilon_{z,t-1} < 0\} \tag{5}$$

where $S_{z,t}^V$ is a latent regime variable governing the volatility state. The vector $Q_{z,t-1}$ contains a number of structural instruments that may affect the level of the conditional asset-specific volatility. $I\{\varepsilon_{z,t-1} < 0\}$ is an indicator function that takes on the value of 1 when $\varepsilon_{z,t-1} < 0$ and zero otherwise. In the case of one regime and $\varkappa = 0$, this model collapses to the asymmetric GARCH model of Glosten, Jagannathan, and Runkle (1993). Similarly, a regime-switching GARCH model is obtained when $\varkappa = \psi_{z,3} = 0$. If one furthermore assumes that $\psi_{z,1} = \psi_{z,2} = 0$, the model reduces to a regime-switching normal

[7]Bekaert et al. (2005b) show that a specification with both a global and regional factor and time-varying factor exposures adequately models the ex-post covariance (correlation) structure of a large set of country-industry and country-style portfolios. Interestingly, this model only slightly underperforms APT specifications, while it strongly outperforms the dummy variable model of Heston and Rouwenhorst (1994) and specifications with constant factor exposures. In Section IV.C, we will nevertheless test whether our specification adequately captures the covariance between the regional, country, and industry indices.

model. Notice that a (regime-switching) (asymmetric) GARCH model is inherently a stationary model. By including structural instruments $Q_{z,t-1}$ in the variance specifications, we allow for structural changes in an otherwise stationary conditional volatility model. This constitutes an important part of our model, as it allows for a structural change in the exposure to systematic risk to be associated with a change in the level and dynamics of idiosyncratic risk. This additional channel is generally omitted in the (volatility spillover) literature.

We choose for a regime-switching volatility specification for two reasons. First, regime-switching volatility models are better suited for dealing with spurious persistence often observed in GARCH estimates (see e.g. Lamoureux and Lastrapes (1990), Hamilton and Susmel (1994), and Cai (1994)). Second, regime-switching models typically accommodate some of the nonlinearities that may show up in higher order moments, such as skewness and kurtosis, as well as asymmetric volatility (see e.g. Perez-Quiros and Timmermann (2001)).

A correct identification of the various shocks also requires an appropriate specification of the expected global market, industry, regional and country returns. Given the focus of this paper on second moments, we do not explore the complex implications of our factor model for expected returns[8]. As a reasonable alternative, we propose the following expected return specification:

$$\mu_{z,t-1} = \gamma_{z,0} + \gamma_z Z_{t-1} \tag{6}$$

where Z_{t-1} represents a vector of information variables part of the information set Ω_{t-1} that have been shown to predict equity returns. To accommodate for potentially partial equity market integration, we include both global and local information variables.

B Estimation Procedure

To keep estimation feasible, we use a three step procedure. First, we estimate the global market shocks. Second, we relate the different regional and global industry returns to the market shocks obtained in the first step. To keep the estimation tractable, we estimate all specifications region (industry) by region (industry). Third, we relate country shocks

[8]See Bekaert and Harvey (1995), De Santis and Gerard (1997), and Carrieri, Errunza, and Sarkissian (2004a) for models exploring the implications of partial integration on expected returns.

to both world and regional shocks. As in the second step, we estimate the specification for each country individually. All estimates are obtained by maximum likelihood. While we report QML standard errors, we do not correct for sampling error of the global (regional) market model parameters in the first (second) stage estimation. Consequently, this approach yields consistent but not necessarily efficient estimates.

A first important assumption[9] behind this three-step procedure is that region and industry-specific shocks are independent from global market shocks and that country-specific shocks are independent from both region-specific and global market shocks. Second, idiosyncratic shocks in one region should be independent from shocks in another region. Third, country and industry-specific shocks should be mutually independent. Recent evidence by Bekaert et al. (2005b) suggests that our specifications with time-varying factor exposures should be sufficiently rich to eliminate most residual asset correlation. We will nevertheless test for residual correlation between the asset-specific shocks further on.

Finally, we need to specify how the underlying states evolve over time. Our most general models feature two latent regime variables, respectively governing time variation in the conditional market betas and volatilities. To limit the number of parameters to be estimated, we make a number of assumptions common to the regime-switching literature. First, we impose the beta and volatility regimes to evolve independently over time. Second, we limit the number of states to two. This assumption has as an additional advantage that it facilitates the identification of the states as business cycle expansions and recessions. Finally, we impose constant transition probabilities. We use the maximum likelihood algorithm first introduced by Hamilton (1989) to estimate the regime-switching beta specifications, and the one of Gray (1996) for models featuring also regime-switching volatilities.

[9] Appendix A in Bekaert and Harvey (1997) provides a formal derivation of all the conditions under which the joint likelihood of a similar (yet less complex) system can be decomposed into a number of univariate models. A similar derivation (but for a two-factor instead of a one-factor model) is available from the working paper version of Baele (2005).

III Data

A Stock Return Data

The dataset consists of weekly US dollar denominated total return indices and market capitalizations for 4 regions, 21 countries, and 18 global industries (see Table 1) over the period January 1973 - August 2007[10]. The US 1-Month Treasury Bill rate is used to calculate excess returns. Our global market portfolio is a value-weighted average of all countries in our sample. Given their size, both the US and Japanese markets are treated as regions. All indices are value-weighted and are obtained from Datastream International. Our sample contains 14 European countries, both from within and outside the EMU, 4 Pacific countries, as well as Canada, Japan, and the US. The industry classification is based on the broad distinction of 18 global industries according to the Dow Jones Indexes and FTSE Industry Classification Benchmark.

A preliminary investigation of the raw returns in Table 2 yields a number of interesting insights[11]. A first observation is that country returns are on average more volatile than industry returns. Not surprisingly, both country and industry returns are considerably more volatile than global market returns. The large difference in volatility between the global market and country portfolios suggests an important role for international diversification in reducing portfolio risk. The relatively smaller difference between global market and industry risks would suggest that country diversification has - at least unconditionally - more potential than industry diversification. Second, we find average intra-industry correlations (63%) to be considerably higher than average country (42%) and regional (42%) correlations, a further confirmation that over the last 30 years, the potential of geographical diversification was on average larger than that of industry diversification. Third, a subsample analysis reveals that cross-country correlations have increased quite substantially over the last three decades, while average cross-industry correlations have remained largely at the same level.

[10]By denominating all returns in US dollar, we take the point of view of the US investor. We are aware that this introduces a common exchange rate component in the asset returns. Griffin and Stulz (2001), however, show that exchange rate shocks are small relative to common market and industry-specific shocks, even for industries that produce internationally traded goods. A number of basic robustness checks indicate that our results are not driven by exchange rate considerations (see Section V.D).

[11]We refer to an appendix available from the authors' websites for detailed summary statistics.

B Structural Instruments

One of the goals of this paper is to investigate to what extent globalization and regional integration have structurally changed the correlation structure of international equity market returns, both across countries (regions) and industries. We allow the (gradual) process of further integration to affect cross-asset correlations by conditioning both the global (regional) market betas and the conditional volatility process on a number of structural economic variables. We focus on two main information variables, namely a trade and an alignment measure. We focus on these two measures because (a) they are theoretically well-founded, (b) they have been successfully used in previous research, (c) they have a high correlation with other potential indicators (such as market development indicators or Quinn (1997)'s integration indicator), and (d) they are both available at the regional, country, and industry level. Finally, we also investigate whether the introduction of the euro in 1999 has had an additional effect on market betas and volatilities.

1 Trade Integration

At the country level, the trade integration measure is calculated as the ratio of imports plus exports over GDP. The empirical model distinguishes between global and regional market shocks, and so does our trade measure. More specifically, the trade integration measure entering the regional market beta only considers the country's trade with other countries within the region that the country belongs to. Similarly, the trade variable entering the global market beta contains the country's trade with all countries outside its region. All data is quarterly and has been obtained from the OECD[12].

Previous studies have successfully linked similar trade integration indicators to cross-country equity returns. Chen and Zhang (1997) for instance found that countries with heavier bilateral trade with a region also tend to have higher return correlations with that region. Bekaert and Harvey (1997), Ng (2000), Bekaert et al. (2005a), and Baele (2005) found that the exposure of country returns to the global (regional) equity market typically increases with measures of trade integration. Forbes and Chinn (2004) showed that, despite the recent growth in global financial flows, direct trade continues to be the

[12]The Import and Export data are from the module 'Monthly Foreign Trade Statistics' from the OECD. All data is seasonally adjusted and converted to a weekly frequency through interpolation. The data is expressed in US dollar.

most important determinant of cross-country equity market comovements. Frankel and Rose (1998) found that countries with closer trade linkages tend to have more correlated business cycles, which should in turn result in higher correlation between their equity markets as well. Trade integration may also proxy for financial integration, and hence a convergence of cross-country risk premiums. For instance, Bekaert and Harvey (1995) found that countries with open economies are generally better integrated with world capital markets.

This study is the first to our knowledge to investigate the effect of trade openness at the industry level on industry betas[13]. We measure industry trade openness by calculating the ratio of the industry's trade relative to its value added. Both the trade and production data is obtained from the STructural ANalysis (STAN) database of the OECD[14]. Theory gives little guidance on the expected effect of trade openness on industry betas. On the one hand, further trade, especially with other industries, may increase the industry's exposure to global market shocks. For instance, Diermeier and Solnik (2001) found that the sensitivity of firm-level stock returns to global market shocks is positively related to the firms' foreign to total sales ratio. On the other hand, further integration, here instrumented by industry openness, may induce investors to focus more and more on industry-specific factors. The effect of the latter channel on betas is, however, unclear.

Panel A of Figure 2 plots an index of the evolution of trade integration at the regional, country, and industry level. A first observation is that trade integration has increased at all levels of aggregation, and especially from the mid-1990s on. Second, at the country level, trade integration with respect to the world and region increase roughly at the same speed. Since 1973, trade integration has on average doubled. Finally, while the increase is relatively higher at the industry than at the country level, this is likely due to the fact that the industry average is based on trading industries only.

[13] Campa and Fernandes (2006) use a similar measure to explain industry effects within the Heston and Rouwenhorst framework.

[14] Industry trade data is available for traded-goods industries and at yearly frequency only. We transform the trade variable to the weekly frequency by means of interpolation. Traded-goods industries are defined in Table 1.

2 Misalignment

At the regional and country level, equity market returns could deviate because of differences in the indices' industrial composition, as pointed out by e.g. Roll (1992). This means that as the industrial structure of a region or country becomes more aligned to that of another region or country, the returns of the equity portfolios should become more similar. Moreover, as the industrial structure of a particular region or country resembles that of the world portfolio, the equity portfolio of that region or country should behave in a similar way as the world portfolio. This implies that the world beta of regions and countries should converge to levels closer to one as industry misalignment decreases. The misalignment of the industrial composition of regions/countries relative to the world is measured as the square root of the mean squared errors between industry weights, i.e.

$$X^{w}_{reg(c),t} = \sqrt{\sum\nolimits_{i=1}^{N_{reg}} \left(w^{reg(c)}_{i,t} - w^{w}_{i,t}\right)^2}, \tag{7}$$

where N_{reg} is the number of industries, $w^{reg(c)}_{i,t}$ the weight of industry i in region reg (country c) and $w^{w}_{i,t}$ the weight of industry i in the world. Weights are computed as the market capitalization of a certain industry in a particular region (country) to the total market capitalization in that region (country). Market capitalizations are obtained from Datastream International. For countries, we also compute the misalignment of the industrial structure of the country relative to the region it belongs to.

As in Carrieri et al. (2004b), we construct a measure for the (mis)alignment of the regional (country) composition within an industry relative to the regional (country) composition of the world portfolio. An industry that is mainly located in one region (country) is likely to be less affected by world shocks, especially when the particular region (country) only makes up a small part of the world. The regional (country) misalignment measure is computed as in equation 7. We expect the world beta of an industry to be negatively related to the misalignment measure.

Panel B of Figure 2 plots an index of the evolution of industry misalignment at the regional, country, and industry level. For countries, industry misalignment decreases substantially both with respect to global and regional equity markets. The effect is slightly higher with respect to regions (minus about 40 percent) than to the global market (minus about 30 percent). A moderate decrease is observed at the regional level (minus about 15 percent). Surprisingly, country misalignment increases for global industries, even though cross-sectional dispersion between industries is large. We do find evidence of a structural decrease for the sectors Media, Transport, Food Retailers, Telecom, Banks, Real Estate, and Investment Companies, but an increase for Automobiles and Parts, Food Products and Tobacco, General Retailers, and Oil and Gas.

3 Euro Dummy

The European integration process culminated in the introduction of a single currency, the euro, in 1999 in 11 European countries. Later on, several other countries joined the euro area, namely Greece (2001), Slovenia (2007), and Malta and Cyprus (2008). While some European countries, like e.g. Denmark, did not join the euro, their exchange rate is highly correlated with the euro. The introduction of the euro may have affected exposures and correlations in a number of ways. First, with the introduction of euro, all exchange rate risk within the euro area disappeared, which should have resulted in increasing correlations between euro area equity markets. Second, the elimination of exchange rate risk may have stimulated European investors to invest outside their home country, either directly or indirectly through a reduction in asymmetric information. A number of studies, including Baele, Pungulescu, and Ter Horst (2007b) and De Santis and Gerard (2006) suggest that the home bias of European investors indeed decreased substantially after the introduction of the euro. Consequently, the marginal investor will increasingly become a European rather than a local investor, and stocks in different countries will increasingly be priced according to a European-wide discount rate, and will less and less depend on the particular degree of risk appetite of local investors. Third, the euro may have served as a facilitator for other integration-stimulating policy measures, such as for instance improvements in trading and settlement infrastructure and corporate governance and reporting procedures. Last but not least, the euro may also have effects outside the euro area through portfolio rebalancing. The optimal weights of all assets in a global portfolio will obviously move with changing expected returns, volatilities, and correlations in part of the assets. The resulting portfolio rebalancing may consequently change the dynamics of returns on all assets, i.e. also of those outside the euro area. To investigate this, we will not only include a euro dummy for euro area countries only, but also for other countries and global industries. The euro dummy is a vector with zeros before January 1999, and ones thereafter.

IV Estimation results for Structural RS Spillover Model

This section summarizes the main estimation results for the structural regime-switching spillover model outlined in Section II. We differentiate between these general and more restricted specifications using standard specification tests. In this Section, we discuss the estimation results for the preferred models. We refer to an appendix available from the authors' websites for detailed model selection statistics. In the first subsection, we discuss the market beta dynamics of the regions, countries, and global industries, to discuss their volatility dynamics in a second subsection.

A Market Beta Dynamics

Panels A and B of Table 3 report the estimated beta dynamics at the regional and global industry level. Panels C and D show the estimation results at the country level.

We find strong evidence that market betas at all levels are driven both by a latent regime variable and the structural instruments, i.e. we strongly reject the null hypothesis of unit and constant betas. The evidence is particularly overwhelming for the regime-switching feature of market betas, which cannot be rejected[15] for any region, industry, or country. The estimates for P and Q are nearly always close to one, indicating that the persistence in market betas (see e.g. Ghysels and Jacquier (2005)) is to a large extent due to a persistence in regime. Regime-Switching betas are (weakly) correlated with business cycle measures such as the term and credit spread, suggesting that market betas contain a business cycle component. For none of the regions, countries, and industries, the latent regime variable exhibits a break-type behaviour, suggesting that permanent shifts (breaks) are well captured by the structural instruments.

The market betas do not only vary with the latent regime variable, but also with our structural instruments. In fact, our model selection procedure selects models with structural instruments for all regions, 15 out of 18 industries, and all but one country. The trade integration variable has a significantly positive effect on the European and Pacific market betas, and, at the country level, on 10 (9) of the global (regional) market be-

[15]Because of the presence of nuisance parameters under the null of one regime, we cannot apply standard asymptotic theory to test for multiple regimes. We use an empirical Likelihood Ratio test instead. More details can be found in an appendix available from the authors' websites.

tas[16]. Not surprisingly, the effect is particularly strong in Europe, a region that has made considerable progress towards further economic and financial integration. The positive effect of trade on market betas could be the result of a convergence in cash flow shocks through further economic integration, an increase in cross-market participations of firms (emergence of more multinationals), or through a convergence in cross-country discount rates (to the extent that further financial integration is correlated with higher degrees of trade integration). We find that trade has both an effect on global and regional betas, indicating that globalization may be at least as important in this respect as regional integration. Industry trade has a significant effect on the market betas of 6 of the 10 trading industries, even though the effect is only economically relevant for the Oil & Gas, Technology, and Construction & Materials industries.

The second structural instrument, industry misalignment, has the expected negative relationship for all regions, and is strongly significant in Europe and the US. This indicates that betas tend to decrease (increase) when a region becomes increasingly different (similar) in its industrial structure from global markets. Similarly, we find a significantly negative impact of industry misalignment in 10 out of the 19 countries[17]. In Finland and Norway, the relationship is significantly positive, which is likely to be caused by their increasing concentration in high beta industries (respectively Telecom and Basic Resources). Similarly, at the industry level, we find a negative effect of country misalignment on the market betas of 10 of the 18 industries[18], suggesting that industry returns become more and more exposed to global shocks when that industry is more evenly spread across countries.

As argued in Section 3, the introduction of the euro in 1999 may have had an effect on common market exposures, even on non-European countries and global industries. At the regional level, we do not find a significant euro effect. At the country level, we find a significant euro effect on the global market betas of 12 of the 19 countries. Except for Austria and Finland, the euro has resulted in an upward shift in global market betas that is not only statistically but also economically relevant (on average, + 0.14). Surprisingly,

[16]We find a negative trade integration effect for Japan, for the global market beta of Norway, and for the regional market betas of Ireland and Canada.

[17]In case of Belgium and Spain, the effect is only significant at the 10 percent level.

[18]In case of Construction & Materials and Healthcare, the effect is only significant at the 10 percent level.

the effect is on average larger for countries not part of the euro area than for those that are (0.23 versus 0.13). Equally surprising is that there is hardly any effect of the euro on the regional market betas (on average, -0.02). Finally, we find a signficant euro effect for 5 of the 18 industries only. For these industries, the euro effect is quite large in absolute value. Just as for the effects at the regional and country level, it remains to be seen whether these effects are truly euro effects.

The different panels in Table 3 also report both the total beta and its structural component over four subperiods, namely the 70s, 80s, 90s, and 00s. A number of interesting patterns emerge. First, for Europe as a region, we observe a gradual structurally-driven increase in its global market beta. While in the 70's the European global market beta was about 0.16 below its regime-switching component, in the 00s it is about 0.17 above that level, resulting in a total increase in market beta of nearly 0.33. A similarly structurally-driven upward effect is found for the Pacific. Second, we find a comparable increase in the structural component of global and regional market betas for most countries. The increase is substantially larger for the global than for the regional market beta (on average, 0.22 versus 0.13 in absolute value, or with 38 versus 28 percent). Surprisingly, we observe only minor differences between euro area and other European markets. For the UK, we observe a mild increase in the structural component of the global market beta, but a substantial decrease in its European market beta, especially in the last period. Also outside Europe, we observe a structural increase in market betas for most countries. In the case of Australia and Canada, we observe a shift from the regional to the global market beta, while the opposite happens for New-Zealand. Third, the structural instruments also lead to substantial changes in the industry market betas. While the market betas of some industries have gradually increased (e.g. Telecom, Banks, Technology), the betas of other industries have decreased (e.g. Oil and Gas, Personal and Household Goods, Health Care).

While we show that market betas have structurally increased for many regions and countries, value-weighted market betas should add up to one in equilibrium. The complexity of our model does not allow us to incorporate cross-asset restrictions, and hence we risk accumulating estimation error. The ex-post analysis in Panel B of Table 5 reveals, however, that a market-weighted average of betas either across regions, countries, and industries is very close to one (respectively, 0.97, 0.98, and 0.97). Moreover, the value-weighted betas do not exhibit an upward trend, and typically fluctuate in the narrow range of 0.95 - 1.05.

The structurally-driven increase was particularly strong at the regional and country level. A detailed look at our results reveals that the structural increase of the global market betas of most European and Pacific betas towards one is accommodated by a decrease in both the weight and the global market beta of the US. In Europe, the increase in the regional market beta in most countries is compensated by a decrease in the regional market beta of the UK.

B Volatility Dynamics

Panels A, B, and C of Table 4 report the estimation results for the volatility specification at respectively the world/regional, industry, and country level.

Global market volatility is best represented by a two-state regime-switching asymmetric GARCH model (see Panel A of Table 4). A number of features are noteworthy. First, volatility regimes are well identified both statistically and economically. As can be seen from Figure 1, the model always attaches a probablity close to one to either the low or high volatilty regime. Moreover, the parameter estimates imply the level of volatility to be more than two times higher in the high volatility state. Second, both volatility regimes are highly persistent, allowing the GARCH parameter estimate to decrease from 0.88 in the AGARCH model to 0.57 (low volatility regime) and 0.74 (high volatility regime) in the case of a RS-AGARCH model. This suggests that the persistence in stock market volatility is also caused by persistence in the volatility regime and only partly by within-regime volatility persistence. Third, we find substantial differences across regimes in the way the conditional volatility reacts to (negative) shocks. While both the ARCH and asymmetry parameters are insignificant in the low volatility regime, both are strongly significant in the high volatility state. Interestingly, in the high volatility state, conditional volatility increases strongly with negative shocks, but actually decreases in response to positive news. This further underlines the need to allow for multiple regimes in conditional volatility models.

In the previous section, we found a positive effect of trade integration, industry structure alignment, and the introduction of the euro on the market betas of both regions and countries. The results in Table 5 show that the same instruments that lead to an increase in market betas reduce both region and country-specific risk. To our knowledge, we are the first to document this structurally-driven shift from idiosyncratic to systematic volatility.

In the beta specification, both trade integration and industry misalignment effects were statistically and economically signficant. In the volatility equation, most of the action comes from the misalignment measure, which is statistically significant and positive for 3 of the 4 regions, and 11 of the 19 countries. While the trade integration measure nearly always has the expected negative sign, it is also mostly statistically insignificant. Interestingly, volatility is significantly lower after the introduction of the euro in 13 of the 19 countries, and 3 of the 4 regions[19], on average with about 1.3 percent (absolute terms).

The right-hand side of Panels A to C also report the total volatility as well as its structural component over different subperiods. We observe a number of interesting patters in the structural component. First, we find a substantial decrease in the structural component of European and Pacific-specific volatility, respectively with about 3.4 percent and 4.4 percent in absolute value. As a percentage of the average level of region-specific volatility, this corresponds to a decrease with 35 and 28 percent. As for the structural component of market betas, no clear pattern emerges for the US and Japan. Country-specific volatility decreases structurally in all countries except New Zealand. On average, the decrease corresponds to 2.85 percent in absolute value, and about 20 percent as a proportion of average volatility. As for market betas, we do not find significant differences between European and other countries.

While evidence for a structural effect is not as strong for global industry-specific volatility as for regions and countries, we nevertheless find a negative effect of trade and a positive effect of country misalignment on the volatility in respectively 3 and 7 global industries. The euro dummy is significant for three industries, and particularly large and positive for the industry ‘basic industries’. It remains to be seen whether this effect can really be atrributed to the introduction of the euro. On average, we find a decrease in industry-specific volatilty of on average -1.14 percent in absolute value, and of 14 percent as a proportion of average volatility.

C Residual Correlation

As discussed in Section II.B, our three-step estimation procedure requires that the factor models are sufficiently rich to eliminate all residual correlation between the region,

[19]The effect of the euro on region-specific volatility is only signficant at the 10 percent level in Europe and the Pacific.

country, and industry-specific shocks. Panel A of Table 5 reports average residual correlations both within and across regions, countries, and industries. Residual correlations are typically lower than 0.03 in absolute terms, and hence are negligible. We do find some negative correlation though between regions, even though the residual correlation is much lower (in absolute value) than the sample correlation, shown in Table 2. Even for selected subsamples, residual correlations are mosly below 10% in absolute value and are substantial lower than the subsample correlations of the raw returns. Generally, this test suggests that our time-varying factor model does sufficiently well in describing cross-asset correlations. This confirms the findings of Bekaert et al. (2005b), who found that a specification with both a global and regional factor and time-varying factor exposures adequately models the ex-post covariance (correlation) structure of a large set of country-industry and country-style portfolios.

V Implications for Portfolio Diversification

We focus on two indicators to assess the portfolio implications of our models. First, we investigate whether the relative size of average idiosyncratic volatility at the regional, country, and global industry level has changed over time. Investors will want to pursue strategies that maximally reduce their exposure to idiosyncratic risk. A rise in the potential of industry diversification would be consistent with a relative increase in average industry-specific relative to country-specific idiosyncratic volatility. Moreover, we quantify the bias in the measures of average idiosyncratic risk that would be induced by not allowing for structural (cyclical) variation in the exposure to common factors and the level of idiosyncratic risk. Second, we analyze how the correlation structure implied by our model estimates has changed over time. A structural increase in cross-country correlations that is not matched by a similar increase in cross-industry correlations would be a further confirmation of a relative increase in the potential of industry diversification. To close this section, we investigate whether the portfolio implications are robust to a change in currency.

A Evolution of idiosyncratic volatility

We measure average idiosyncratic volatility as follows:

$$\sigma_{Z,t} = \sum\nolimits_{z \in Z} w_z h_{z,t}^{1/2}$$

where w_z and $h_{z,t}$ represent asset z's market weight and idiosyncratic variance at time t, and Z contains all assets over which to aggregate.

Panel A of Figure 3 plots the results of this aggregation at the regional, country, and industry level. The shaded areas represent global recession periods. We find a number of interesting results. First, even after correcting for structural and cyclical variation in market betas, our results indicate that average idiosyncratic risk both across countries (regions) and industries shows a strongly countercyclical pattern. This suggests that, despite the well-known asymmetry in equity correlations (see e.g. Ang and Bekaert (2002)), international diversification continues to pay off in times of recession and/or market turmoil. Second, we find that a considerable part of country-specific risk can be eliminated by diversifying regionally. However, diversifying not only across regions but also across countries results in non-negligible further risk-reduction benefits. Third, unlike Griffin and Karolyi (1998), we find only a small difference in average industry-specific risk when we aggregate over respectively 10 and 18 industries. Fourth, over the period 1973-1999, industry-specific volatility is consistently lower than both region- and country-specific volatility. This confirms findings in previous papers that during this period investors were better off diversifying their portfolios across countries rather than across industries. Notice, however, that the edge of average country-specific over average industry specific risk is at times limited (especially in the second half of the 1980s) and often much smaller than found in unit/constant beta studies (see further). Fifth, at the end of the 1990s, we observe a strong rise in the level of industry-specific risk, an increase that is only partially matched by an increase in country-specific risk. In 1999, industry-specific risk surpassed country-specific risk for the first time in nearly 30 years, to peak at the end of 2001. Based on similar observations, previous studies concluded that sector diversification strategies had become at least as beneficial as geographical diversification strategies. We come to a different conclusion. Industry-specific risk declined substantially from the end of 2002 on, to levels below average country-specific risk from 2003 onwards. This suggest that previous studies have called too quickly for the end of the dominance of geographical over industry diversification.

Brooks and Del Negro (2004) argued that the relative increase in industry risk at the end of the 1990s may have been a purely temporary phenomenon related to the TMT bubble. To analyze this, we calculate the average industry-specific risk taking into account all but the TMT industries[20]. To fully eliminate the effect of the TMT bubble, we remove the TMT industries from the global and all regional and country indices, and re-estimate the optimal volatility spillover models[21]. In Panel B of Figure 3, we plot average region, country, and industry-specific volatility excluding the TMT industries over time. While the level of average region and country-specific volatility is relatively unaffected, excluding the TMT industries leads to a substantial decrease in the level of industry-specific risk at the end of the 1990s. Interestingly, even after excluding the TMT industries, we still find a significant increase in industry-specific risk, suggesting that the TMT bubble was only partially responsible for the surge in industry risk at the end of the 1990s.

Finally, we investigate to what extent the evolution of country-specific risk is different for Europe. Given that over the last 20 years Europe has gone through an extraordinary period of economic, monetary, and financial integration, it is not surprising that structural changes both in the market beta and idiosyncratic volatility were most apparent in Europe. Figure 4 plots the different components of total risk for the (weighted) average European country. We find a number of interesting patterns. First, we observe a very clear downward trend in average idiosyncratic volatility. This decrease is substantial in economic terms, from about 15% in the early 1970s to about 10% in the more recent period, or a decrease by more than 30 percent. To our knowledge, this study is the first to document and quantify such a decrease. Second, we find a similar yet slightly less outspoken downward trend in the average regional risk component across countries. This is mainly the result of a gradual reduction in the level of European-specific risk resulting from an increased exposure of the aggregate European market to global market shocks. Third, we observe a modest increase in the importance of global market risk over the period 1973-1996. The market component increases substantially, though, during the 1997-2000 period, reaching an all-time peak in April 2001. While the market component has decreased since then, it continues being higher than at any moment in the previous three decades.

[20] More specifically, we remove the Telecom, Media, IT Hardware, and IT Software industries.

[21] We obtain very similar results when we do not eliminate the TMT sectors from the global and regional benchmark indices, i.e. when we allow for spillovers from the TMT to other industries.

B Unit (constant) betas and the bias in idiosyncratic risk

In the introduction, we argued that the assumption of unit (constant) betas typically made in the country-industry literature is not only likely to be rejected by the data, but that it may also lead to a substantial mismeasurement or 'bias' in the potential of geographical and industry diversification strategies. In this section, we first quantify the bias in measures of average region, country, and industry-specific risk induced by not allowing betas to be constant (instead of unity) or time-varying. Second, we investigate what extensions of the unit beta model are most important, i.e. are crucial in reducing the total bias. This should help future studies decide about the optimal level of model complexity.

The biases induced by assuming unit betas in case of a one-factor (for regions, industries) and a two-factor (for countries) model are given respectively by

$$bias^1_{Z,t} = \frac{\left(\sum_{z\in Z} w_{z,t}(\beta^w_{z,t}-1)^2\right) h_{w,t}}{\sum_{z\in Z} w_{z,t} h_{z,t}} \quad . \tag{8}$$

and

$$bias^1_{C,t} = \frac{\left(\sum_{c\in C} w_{c,t}(\beta^{reg(c)}_{c,t}-1)^2\right) h_{reg(c),t} + \left(\sum_{c\in C} w_{c,t}\left[(\beta^w_{c,t}-1)-(\beta^w_{reg(c),t}-1)\right]^2\right) h_{w,t}}{\sum_{c\in C} w_{c,t} h_{c,t}} \quad . \tag{9}$$

where Z contains either the regions or the global industries, and C the countries over which to aggregate. $reg(c)$ refers to the regional market reg the country c belongs to[22]. The biases induced by constant betas can be derived in a similar way. Equation (8) indicates that measures of average region (industry)-specific volatility using unit market betas are positively biased relative to our measure by the average cross-sectional variance in the betas (relative to unit betas) times the conditional world market variance. Similarly, equation (9) shows that unit beta models typically overestimate average country-specific volatility by an amount that is positively related to first the average cross-sectional variance of the country's global market exposure relative to the region's global market exposure times the world variance and second by the average cross-sectional variance of the region's global market exposure relative to unit beta case times the region-specific variance. Notice that the country bias reflects the bias in country-specific risk only, excluding the

[22] We refer to an appendix available from the authors' websites for the full derivation of the biases.

bias in region-specific risk. To make the biases in industry and country diversification potential comparable, however, the region and country-specific biases are aggregated[23].

Figure 5 plots the bias in the measures of average idiosyncratic risk resulting from imposing respectively unit (Panel A) and constant betas (Panel B). A number of interesting patterns emerge. First, assuming unit betas creates a potentially large bias in measures of both industry and country risk. Country-specific risk is overestimated by nearly more than 25 percent in the early 1970s and by about 15 percent in the period 1985-1995. The bias in average industry-specific risk is generally below 10 percent, except during the years corresponding to the TMT bubble when it increases to nearly 30 percent. In absolute terms, this means that average idiosyncratic volatility is overestimated by up to four (countries) and five (industries) percent. The size and timing of this bias sheds some light on the findings of studies that use unit beta models. On the one hand, our results suggest that the strong outperformance of country over industry diversification strategies found until the mid-1990s has been overstated. On the other hand, our findings also indicate that the recent surge in industry risks has been seriously overstated. Consequently, unit beta models seem to overstate the relative change in geographical and industry diversification at the end of the 1990s. Second, the bias in average industry-specific risk decreases substantially when the TMT sectors are not taken into account, from nearly 30 percent to less than 20 percent. This is easily understood by observing that the betas of especially the TMT sectors during this period were considerably above one and hence contributed substantially to the first term in the bias, $\left(\sum_{z\in Z} w_{z,t}(\beta_{z,t}^{w}-1)^2\right)$. Third, as can be seen from Panel B, the bias at the country level does not decrease substantially when constant instead of unit betas are allowed for, further underlining the need for time-varying betas at the country level. At the industry level, however, a considerable part of the bias disappears when betas are allowed to be different from one but constant, except during the TMT bubble when the bias remains considerable (about 17 (11) percent including (excluding) the TMT industries). We do observe, though, a small increase in the bias in years with large positive returns, providing indirect evidence in favour of Santos and Veronesi (2004)'s hypothesis that the cross-sectional dispersion in betas increases when the market risk premium decreases.

Finally, we investigate what features of our model are most important for reducing the to-

[23] The appendix available from the authors' websites shows how this can be accomplished.

tal bias. We respectively quantify the contribution of allowing betas to be constant instead of being unity, of allowing structural instruments in the betas (relative to the constant beta case), of allowing regime-switches in the betas (relative to the beta specification with instruments), and finally of also allowing for structural shifts in the asset-specific volatility specification (relative to model with time-varying betas and an AGARCH volatility specification). The individual contributions sum up to the total bias. Figure 6 reports the decomposition at the country level (Panel A) and at the industry level (Panel B). The decomposition yields a number of interesting insights, and confirms the findings from the previous paragraph. First, the bias in industry-specific risk is reduced considerably when constant instead of unit market betas are allowed for. Not surprisingly, the exception is the period corresponding to the TMT bubble, during which especially the regime-switching component contributes to a reduction in the total bias. Second, the bias in total country-specific risk reduces only marginally when betas are allowed to be constant instead of unity. Both instruments and the latent regime variable are required to reduce the bias to zero. Finally, neither the euro nor the instruments in the specification for asset-specific volatility contribute much to a reduction in the total bias.

C Evolution of model-implied correlations

In this section, we investigate the evolution of average conditional correlations over time, both across countries (regions) and industries. We focus on three questions. First, we investigate to what extent correlations are asymmetric, i.e. higher in highly volatile periods. Second, we analyze whether further integration and globalization have lead to a structural increase in cross-country correlations. Finally, we compare the relative size of cross-country and cross-industry correlations over time. A structural increase in cross-country correlations that is not matched by a similar increase in cross-industry correlations would be consistent with a decrease (increase) in the potential of geographical (industry) diversification.

Assume that the asset-specific shocks e_t are uncorrelated. Under this assumption, the conditional correlation between two regions or industries i and j is given by

$$\rho_{i,j,t} = \beta_{i,t}^{w}\beta_{j,t}^{w}\frac{\sqrt{h_{w,t}}}{\sqrt{\left(\beta_{i,t}^{w}\right)^{2}h_{w,t}+h_{i,t}}}\frac{\sqrt{h_{w,t}}}{\sqrt{\left(\beta_{j,t}^{w}\right)^{2}h_{w,t}+h_{j,t}}} \tag{10}$$

where the symbols are defined as before. Given the substantial evidence in this and previous papers of no trend in global equity market volatility, equation (10) clearly shows that a *structural* increase in cross-asset correlations is the result of a structural increase in the assets' market beta and/or a fundamental decrease in the level of asset-specific risk. A similar formula can be derived for the cross-country correlations, which will in addition be driven by the time-varying regional market betas as well as the regional market's volatility. We calculate average market-weighted correlations as follows:

$$\bar{\rho} = \frac{\sum_{i \in Z} \sum_{j \in Z, j \neq i} \omega_i \omega_j \rho_{i,j}}{\sum_{i \in Z} \sum_{j \in Z, j \neq i} \omega_i \omega_j} \tag{11}$$

where Z contains respectively all regions, countries, or global industries.

Panel A of Figure 7 plots the evolution over time of the average market-weighted conditional correlations across regions, countries, and industries. A number of interesting patterns emerge. First, we observe a clearly cyclical pattern in the average correlations for all asset classes: correlations are on average substantially higher in recessions than in expansions, i.e. asymmetric. Second, there is a strong upward trend in average cross-regional and cross-country correlations, indicating a reduction in the benefits from international diversification. In the case of countries, average correlations were typically below 40 percent in the 1970 compared to up to 65 percent in the recent years. The increase is even more substantial (up to nearly 85 percent) when we look specifically at the continental European equity markets. This suggests that globalization and regional integration both contribute to increasing international correlations. Third, for nearly the entire sample, cross-industry correlations are substantially higher than cross-country correlations, confirming the superiority of geographical relative to industry diversification strategies. Fourth, the period around the TMT bubble had a strong but temporary effect on both average industry and country (regional) correlations. The surge in industry-specific risk at the end of the 1999s led to a substantial decrease in cross-industry correlations, only partially matched by a decrease in cross-country correlations[24]. As a result, from 2000 up to 2004, industry correlations were for the first time in nearly 30 years lower than country correlations. Since 2004, correlations are again larger at the country than at the industry level, even though the margin is much smaller than before.

[24]The decrease in correlations is much smaller at the country than at the industry level because a large part of the surge in industry-specific risk is diversified away at the country level.

Panel B of Figure 7 plots the bias in correlations induced by assuming constant betas. While the bias is rather small at the industry level, it is substantial at the country level. Constant beta models overestimated average cross-country correlations by over 30 percent in early 1970s. While the bias is on average lower in 1980s and early 1990s, correlations are especially overstated in recession periods. The bias becomes substantially negative as from the end of the 1990s on. At the end of the sample, our results imply that constant beta models would underestimate cross-country correlations by more than 20 percent. In unreported results, we found that the bias is even higher for unit beta models. At the industry level, the bias is consistently around 10 percent. At the country level, the bias amounts to nearly 60 percent in the 1970s, to gradually decrease to levels close to zero at the end of the sample.

D Robustness for Alternative Currency Denomination

The results of this paper are based on US dollar denominated returns. We investigate the robustness of our results by estimating the various models and diversification potential indicators using deutschmark (before January 1999) and euro (after January 1999) denominated returns. Our results[25] remain remarkably robust to the change in currency. Industry-specific volatility is consistently lower than both region- and country-specific volatility, with the exception of a short period around 2000. Contrary to the results based on US dollar denominated returns, industry-specific risk surpassed country-specific risk only slighly at the end of the 1990s, to revert to levels below country-specific risk from 2001 onwards. Similarly, for nearly the entire sample, cross-industry correlations were substantially higher than cross-country correlations, as in the US dollar case. Likewise, we notice a strong convergence in the average correlations towards the end of the 1990s. We see furthermore a remarkable increase in average correlations for European equity markets, although levels are consistently lower than for the US dollar case. This is due to the fact that the common exchange rate component in European equity market returns is removed by expressing the returns in deutschmark/euro.

[25] Detailed estimation results are available from the authors.

VI Conclusion

In this paper, we investigate the dynamics and determinants of the relative potential of geographical and industry diversification over the last 30 years. We start by arguing that structural changes in the economic and financial environment, such as time-varying integration and globalization, are likely to be important drivers of the relative benefits of country and industry diversification. While the recent literature has indeed pointed to increasing integration as a possible explanation for the recent shift from country to industry diversification, this study is to our knowledge the first to explicitly embed these structural changes in a fully conditional model for country and industry returns. We estimate a dynamic factor model on a set of 4 regions, 21 countries, and 18 global industries over the period 1973-2007. We relax the common assumption of unit or constant factor exposures (see e.g. Heston and Rouwenhorst (1994), Campbell et al. (2001), Brooks and Del Negro (2002), and Ferreira and Gama (2005)) to the case where these exposures vary with two structural instruments – reflecting time-varying integration and market development – as well as with a latent regime variable – reflecting temporary economic fluctuations. In a similar vein, we also allow our two structural instruments to interact with the level of conditional volatility at the global, regional, country, and industry level. Based upon the model estimates, we calculate and compare two popular indicators of diversification potential both for countries and industries, namely average asset-specific volatilities and model-implied correlations.

We find that market exposures of both countries and industries deviate substantially and for long periods from both one and their unconditional value. The dynamics of their time variation differs, however, substantially between countries and industries. While both the global and regional betas of (mainly European) countries show an important structural increase, the time variation in global industry betas is mainly driven by cyclical factors. The effect of globalization and integration on market betas is not only statistically but also economically important, and amounts to on average 38 and 28 percent for global and regional market betas. Next, we show that this increased market exposure tends to be compensated by a structural decrease in country-specific risk (on average, by about 20 percent of total country-specific risk), keeping total country risk roughly at the same level. This gradual shift from country-specific to common risk is intensified further by the introduction of the euro. Consequently, we demonstrate that not accounting for this time variation in market betas and volatilities leads to substantial biases in measures of both industry and country risk. In our sample, we find biases in average country and industry-specific risk of more than 25 percent. The bias in industry-specific risk is generally below 10 percent, but rises to nearly 30 percent in the period corresponding to the TMT bubble. We show that unit and constant beta studies have overstated the benefits of geographical diversification especially in the early 1970s and between 1985-1995, while overstating the benefits on industry diversification at the end of the 1990s. After correcting for these biases, we find that over the last 30 years average country-specific risk was typically higher than average industry-specific risk. When time-varying betas are accounted for, the edge of geographical over industry diversification is, however, substantially lower than in unit/constant beta studies. Similarly, while cross-industry correlations have typically been above cross-coutnry correlation, the difference has become smaller as integration improved. Similar to other studies, we find a substantial increase in industry-specific relative to country-specific risk at the end of the 1990s, and a corresponding decrease in cross-industry relative to cross-country correlations. Contrary to Brooks and Del Negro (2004), we find that this rise is not a pure artifact of the buildup and burst of the TMT bubble. From about 2003 on, however, cross-industry correlations rose again above cross-country correlations, while average country-specific volatility again surpassed average cross-inudstry volatility. We conclude that while the benefits of geographical diversification have gradually decreased with globalization and integration, they are still substantial, to the extent that geographical diversification yields still larger risk reduction benefits than industry diversification.

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Table 1: Regions, Countries and Industries

The region Europe is disaggregated over 14 European countries. The region Pacific consists of 4 countries (Australia, Hong Kong, New Zealand and Singapore). This results in 21 geographical entities. The industry classification is based on the broad distinction of 18 global industries according to the Dow Jones Indexes and FTSE Industry Classification Benchmark. The traded-goods industries are marked with a 'T', the non-traded-goods industries with a 'NT'.

Region	**Codes**	**Region**	**Codes**
Europe	EU	Japan	JP
Pacific	PC	United States	US
Country	**Codes**	**Country**	**Codes**
Australia	AU	Italy	IT
Austria	OE	Netherlands	NL
Belgium	BG	New Zealand	NZ
Canada	CN	Norway	NW
Denmark	DK	Singapore	SG
Finland	FN	Spain	ES
France	FR	Sweden	SD
Germany	BD	Switzerland	SW
Hong Kong	HK	UK	UK
Ireland	IR	**World**	**WD**
Industry	**Codes**	**Industry**	**Codes**
Oil & Gas (T)	OILGS	Retail (NT)	RTAIL
Chemicals (T)	CHMCL	Media (NT)	MEDIA
Basic Resources (T)	BRESR	Travel & Leisure (NT)	TRLES
Construction & Materials (T)	CNSTM	Telecommunications (NT)	TELCM
Industrial Goods & Services (T)	INDGS	Utilities (NT)	UTILS
Automobiles & Parts (T)	AUTMB	Banks (NT)	BANKS
Food & Beverage (T)	FDBEV	Insurance (NT)	INSUR
Personal & Household Goods (T)	PERHH	Financial Services (NT)	FINSV
Health Care (T)	HLTHC	Technology (T)	TECNO

Table 2: Summary Statistics

All data are weekly US denominated total returns over the period January 1973 - August 2007, for a total of 1807 observations. The returns are in excess of the U.S. one-month Treasury-bill rate. Country portfolio returns are computed as the value-weighted averages of the local industry portfolio returns. Region portfolio returns are computed by aggregating over the relevant countries in the dataset. World portfolio returns are computed by aggregating over the 4 regions. The industry portfolio returns are computed as the value-weighted average of country industry portfolio returns. Panel A shows the average mean and standard deviation over the different sets of portfolio returns, as well as the average standard deviations over different subsamples. The mean and standard deviation values are annualized. The cross-sectional variation is shown between brackets. Panel B displays the average unconditional correlations between the different sets of portfolio returns, as well as the average unconditional correlations between respectively region returns, industry returns and country returns over different subsamples. Individual correlations are averaged using the market capitalizations as weights.

Panel A

A. Means and (Subsample) Standard Deviations								
	Mean		**Stdev**		**73-79**	**80-89**	**90-99**	**00-07**
World	5.59	-	13.38	-	12.61	13.48	12.68	14.73
Region	6.54	*(1.34)*	17.78	*(2.78)*	17.62	18.49	16.92	17.54
Country	8.80	*(2.21)*	21.01	*(4.19)*	21.17	22.04	19.22	19.87
Industry	6.09	*(1.02)*	16.42	*(2.41)*	15.72	16.65	15.11	17.84
B. Unconditional Correlations								
	World	**Region**	**Country**	**Industry**	**73-79**	**80-89**	**90-99**	**00-07**
World	1	0.71	0.53	0.80	-	-	-	-
Region	0.71	0.42	0.44	057	0.35	0.52	0.45	0.61
Country	0.53	0.44	0.42	0.43	0.29	0.36	0.44	0.60
Industry	0.80	0.57	0.43	0.63	0.64	0.65	0.66	0.63

Table 3: Beta Specification and Significance Tests

This table reports the results for the beta specification of the selected model. Panel A shows the results for the regions, Panel B for the industries, Panel C and Panel D for the countries. The column 'Model' shows whether the specifications for the market betas ('β') and conditional variances ('σ') contain a RS component and/or Instruments ('I') and/or a Euro dummy ('E'). The next two colums show the betas over the two different regimes. The p-value of the (Wald) test whether betas are significantly different across regimes is reported between brackets. The columns 'Trade', 'Align' and 'Euro' show the coefficients of respectively the two structural instruments and the euro dummy in the beta specification (if applicable). P-values are reported between brackets. The next two columns report the transition probabilities with their p-values between brackets. The last part of the table shows the total beta (first line) and its structural component (second line) over different subperiods. *, †, and ‡ indicate significance at respectively the 1, 5, and 10 percent level for the t-test whether subperiod betas are equal to the full-period betas (first line) and the t-test whether the structural component is equal to zero (second line, italics). The tests are corrected for autocorrelation. Note that for the countries, we make a distinction between the world market beta and the regional market beta of the country.

Panel A

Region	**Model**	β_1	β_2	**Trade**	**Align**	**Euro**	**P**	**Q**	**73-79**	**80-89**	**90-99**	**00-07**
Europe	β(RS,I),σ(I,E)	1.032	0.673	0.111	-0.055	-	0.969	0.968	0.659*	0.821*	0.852*	1.027*
		(0.00)		(0.002)	(0.018)		(0.000)	(0.000)	*-0.155**	*-0.049**	*0.027**	*0.171**
Pacific	β(RS,I),σ(I,E)	0.967	0.604	0.124	-0.027	-	0.984	0.988	0.672*	0.729*	0.759	0.866*
		(0.00)		(0.000)	(0.198)		(0.001)	(0.000)	*-0.144**	*-0.095**	*0.077**	*0.154**
Japan	β(RS,I),σ(I,E)	1.377	0.618	-0.221	-0.005	-	0.966	0.971	0.713*	1.066*	1.203*	0.904*
		(0.00)		(0.000)	(0.398)		(0.000)	(0.000)	*-0.064**	*-0.061**	*0.203**	*-0.128**
US	β(RS,I),σ(I,E)	1.163	0.812	-0.018	-0.081	-	0.979	0.977	1.161*	0.948*	0.913*	0.958*
		(0.00)		(0.333)	(0.003)		(0.000)	(0.000)	*0.113**	*-0.004*	*-0.067**	*-0.010*‡

Panel B

Industry	Model	β_1	β_2	Trade	Align	Euro	P	Q	73-79	80-89	90-99	00-07
OILGS	β(RS,I),σ()	1.005	0.725	0.233	-0.051		0.982	0.987	1.012	0.933*	0.637*	0.872*
		(0.001)		(0.008)	(0.048)		(0.000)	(0.000)	0.173*	0.072*	-0.238*	-0.131*
CHMCL	β(RS),σ()	0.839	1.095				0.966	0.987	0.989*	1.028*	1.024*	1.035*
		(0.000)					(0.000)	(0.000)				
BRESR	β(RS,E),σ(I,E)	1.212	0.728			0.211	0.994	0.988	1.070*	1.132*	1.006*	1.172*
		(0.000)				(0.070)	(0.000)	(0.000)				0.211
CNSTM	β(RS,I),σ(I)	0.704	1.171	0.125	-0.175		0.956	0.953	0.594*	0.997*	1.002*	1.034*
		(0.000)		(0.085)	(0.100)		(0.000)	(0.038)	-0.337*	0.054*	0.061*	0.059*
INDGS	β(RS,I),σ(I,E)	1.101	0.915	0.040	0.007		0.988	0.978	0.991*	1.003*	1.039*	1.080*
		(0.000)		(0.036)	(0.322)		(0.000)	(0.000)	-0.038*	-0.033*	0.016*	0.047*
AUTMB	β(RS,I),σ()	0.698	1.121	-0.026	0.075		0.972	0.987	1.015*	0.970*	0.974*	1.015*
		(0.000)		(0.300)	(0.008)		(0.022)	(0.028)	0.127*	0.062*	-0.041*	0.073*
FDBEV	β(RS,I,E),σ(E)	0.940	0.740	0.034	-0.019	-0.281	0.996	0.991	0.901*	0.878*	0.798*	0.579*
		(0.000)		(0.044)	(0.075)	(0.000)	(0.011)	(0.005)	-0.018*	0.011*	-0.036*	-0.267*
PERHH	β(RS,I,E),σ(I,E)	0.798	1.039	-0.081	-0.070	-0.092	0.985	0.996	1.137*	0.989*	0.922*	0.742*
		(0.000)		(0.001)	(0.000)	(0.042)	(0.000)	(0.000)	0.123*	0.045*	-0.059*	-0.199*
HLTHC	β(RS,I,E),σ(I,E)	0.862	1.081	-0.066	0.016	-0.315	0.997	0.997	1.119*	0.941*	0.882*	0.583*
		(0.000)		(0.002)	(0.081)	(0.000)	(0.000)	(0.000)	0.067*	0.034*	-0.031*	-0.422*
RTAIL	β(RS,I),σ()	1.241	0.911		-0.042		0.978	0.990	1.123*	1.039*	0.988*	0.955*
		(0.000)			(0.047)		(0.000)	(0.000)	0.043*	0.002	0.000	-0.057*
MEDIA	β(RS,I),σ(I)	1.305	0.911		0.087		0.993	0.999	1.250*	0.980*	0.904*	1.072*
		(0.000)			(0.000)		(0.000)	(0.000)	0.104*	0.055*	-0.045*	-0.062*
TRLES	β(RS,I),σ(I,E)	0.881	1.445		-0.069		0.992	0.982	1.269*	0.924*	0.980*	1.107*
		(0.000)			(0.001)		(0.000)	(0.000)	-0.005	-0.043*	0.025*	0.067*
TELCM	β(RS,I,E),σ(I,E)	1.042	0.663		-0.110	0.251	0.867	0.926	0.697*	0.730*	0.934*	1.090*
		(0.001)			(0.000)	(0.000)	(0.148)	(0.000)	-0.095*	-0.072	0.136*	0.286*
UTILS	β(RS,I),σ(I,E)	0.476	0.815		0.099		0.989	0.997	0.645*	0.763*	0.644*	0.619*
		(0.000)			(0.000)		(0.000)	(0.001)	-0.016*	0.083*	-0.089*	-0.081*
BANKS	β(RS,I),σ(I)	0.783	1.117		-0.145		0.996	0.993	0.651*	0.826*	1.117*	1.028*
		(0.000)			(0.000)		(0.000)	(0.000)	-0.177*	-0.076*	0.060*	0.185*
INSUR	β(RS),σ()	0.926	1.232				0.998	0.988	0.940*	0.960*	0.982*	0.998*
		(0.000)					(0.000)	(0.012)				
FINSV	β(RS,I),σ(I)	1.415	1.071		-0.035		0.991	0.997	1.063*	1.208*	1.293*	1.134*
		(0.000)			(0.109)		(0.000)	(0.000)	-0.045*	-0.029*	0.013*	0.046*
TECNO	β(RS,I,E),σ(E)	1.456	0.966	0.198	-0.003	-0.252	0.999	0.997	1.214*	1.174*	1.227*	1.352*
		(0.000)		(0.000)	(0.397)	(0.005)	(0.003)	(0.000)	-0.232*	-0.124*	0.054*	0.017*

Panel C

Country	Model	Global Factor					Regional Factor				
		β_1	β_2	**Trade**	**Align**	**Euro**	β_1	β_2	**Trade**	**Align**	**Euro**
Austria	β(RS,I,E),σ(E)	1.261	0.575	0.205	-0.083	-0.101	1.367	0.677	0.187	-0.014	-0.188
		(0.000)		(0.000)	(0.042)	(0.000)	(0.009)		(0.000)	(0.388)	(0.033)
Belgium	β(RS,I,E),σ(I,E)	0.404	0.694	0.185	-0.036	0.417	0.573	0.861	0.190	-0.017	0.225
		(0.000)		(0.000)	(0.087)	(0.000)	(0.012)		(0.000)	(0.345)	(0.094)
Denmark	β(RS,I),σ(I,E)	0.487	0.757	0.236	-0.037		0.351	0.872	0.137	-0.156	
		(0.031)		(0.000)	(0.283)		(0.000)		(0.013)	(0.011)	
Finland	β(RS,I,E),σ(I,E)	1.258	1.066	0.250	0.195	-0.359	0.822	4.509	0.144	-0.125	-0.002
		(0.800)		(0.258)	(0.012)	(0.107)	(0.057)		(0.340)	(0.153)	(0.399)
France	β(RS,I,E),σ(I,E)	0.912	0.673	-0.035	-0.066	0.232	1.107	0.757	0.128	-0.003	-0.150
		(0.000)		(0.122)	(0.032)	(0.000)	(0.000)		(0.032)	(0.397)	(0.183)
Germany	β(RS,I,E),σ(I,E)	0.853	0.523	0.165	-0.074	0.325	1.077	0.653	0.079	-0.003	-0.043
		(0.000)		(0.000)	(0.001)	(0.001)	(0.000)		(0.010)	(0.397)	(0.375)
Ireland	β(RS,I,E),σ(I,E)	0.651	1.204	0.023	-0.085	0.250	0.972	-0.397	-0.144	0.014	-0.112
		(0.077)		(0.326)	(0.049)	(0.002)	(0.114)		(0.012)	(0.387)	(0.246)
Italy	β(RS,I,E),σ(I,E)	0.614	0.741	0.040	-0.160	0.147	0.573	0.976	0.088	-0.122	0.205
		(0.008)		(0.222)	(0.000)	(0.065)	(0.000)		(0.086)	(0.019)	(0.236)
Netherlands	β(RS,I,E),σ(I,E)	1.020	0.648	0.093	-0.005	0.167	0.861	0.782	0.034	-0.090	-0.034
		(0.000)		(0.004)	(0.393)	(0.085)	(0.372)		(0.178)	(0.002)	(0.381)
Norway	β(RS,I,E),σ(I,E)	0.432	1.054	-0.064	0.186	0.178	0.521	1.140	0.124	0.097	0.160
		(0.000)		(0.027)	(0.036)	(0.159)	(0.000)		(0.087)	(0.183)	(0.223)
Spain	β(RS,I),σ(E)	1.053	0.986	0.142	-0.060		1.130	0.740	0.168	-0.006	
		(0.211)		(0.039)	(0.098)		(0.018)		(0.044)	(0.199)	
Sweden	β(RS,I,E),σ(I,E)	0.983	0.469	0.172	-0.185	0.331	0.741	3.654	0.111	-0.201	0.082
		(0.145)		(0.000)	(0.000)	(0.019)	(0.000)		(0.059)	(0.001)	(0.360)
Switzerland	β(RS,I),σ(I,E)	0.881	0.568	0.108	0.003		0.833	0.543	0.135	0.023	
		(0.000)		(0.000)	(0.397)		(0.000)		(0.001)	(0.321)	
UK	β(RS,I,E),σ(I,E)	0.806	1.205	-0.001	-0.059	0.187	1.125	1.883	-0.012	0.161	-0.284
		(0.001)		(0.399)	(0.007)	(0.003)	(0.000)		(0.368)	(0.017)	(0.000)
Australia	β(RS,I),σ(I)	0.612	0.719	0.097	0.027		0.546	1.238	0.005	0.042	
		(0.043)		(0.000)	(0.252)		(0.000)		(0.394)	(0.174)	
Hong Kong	β(RS,I,E),σ(I,E)	0.656	0.913	-0.039	0.035	0.150	0.696	1.351	0.087	-0.046	-0.180
		(0.000)		(0.273)	(0.196)	(0.019)	(0.000)		(0.106)	(0.157)	(0.212)
New Zealand	β(RS,I,E),σ(I,E)	0.855	0.423	-0.039	0.018	-0.128	0.792	0.155	-0.059	-0.081	0.355
		(0.092)		(0.241)	(0.363)	(0.331)	(0.000)		(0.177)	(0.092)	(0.012)
Singapore	β(RS,I,E),σ(I,E)	0.847	0.536	-0.011	-0.021	0.191	0.893	0.400	-0.008	-0.091	0.147
		(0.002)		(0.388)	(0.349)	(0.019)	(0.000)		(0.395)	(0.041)	(0.193)
Canada	β(RS,I,E),σ(I,E)	0.543	0.988	0.157	-0.067	0.233	0.459	0.464	-0.157	-0.046	-0.433
		(0.000)		(0.000)	(0.097)	(0.062)	(0.977)		(0.000)	(0.137)	(0.000)

Panel D

Country	Model	Global Factor 73-79	Global Factor 80-89	Global Factor 90-99	Global Factor 00-07	Regional Factor 73-79	Regional Factor 80-89	Regional Factor 90-99	Regional Factor 00-07	P	Q
Austria	β(RS,I,E),σ(E)	0.282*	0.482*	0.628*	0.649*	0.447*	0.587*	0.882*	0.959*	0.994	0.997
		-0.357*	-0.231*	-0.144*	-0.322*	-0.296*	-0.229*	0.007	-0.117*	(0.000)	(0.000)
Belgium	β(RS,I,E),σ(I,E)	0.347*	0.571*	0.631*	0.929*	0.525*	0.625*	0.751†	1.085*	0.988	0.984
		-0.165*	0.071*	0.090*	0.410*	-0.155*	-0.043*	0.042*	0.397*	(0.000)	(0.000)
Denmark	β(RS,I),σ(I,E)	0.420*	0.651*	0.569*	0.874*	0.489*	0.542*	0.790*	0.891*	0.990	0.990
		-0.166*	0.075*	-0.085*	0.257*	-0.054*	0.02*	0.117*	0.290*	(0.000)	(0.000)
Finland	β(RS,I,E),σ(I,E)		0.564*	0.937*	1.414*		0.860*	0.973*	0.957*	0.986	0.679
		0.065*	-0.051*	-0.313*	0.163*	0.027*	0.012*	0.011	-0.005	(0.184)	(0.314)
France	β(RS,I,E),σ(I,E)	0.785*	0.813*	0.882*	1.132*	0.820*	0.916*	1.014†	1.090*	0.992	0.981
		-0.024*	-0.009*	0.037*	0.246*	-0.135*	-0.058*	0.006	0.021*	(0.000)	(0.000)
Germany	β(RS,I,E),σ(I,E)	0.593*	0.653*	0.774*	1.188*	0.713*	0.854*	0.934*	1.024*	0.975	0.963
		-0.058*	-0.031*	0.002	0.458*	-0.104*	-0.007*	-0.038*	0.105*	(0.000)	(0.000)
Ireland	β(RS,I,E),σ(I,E)	0.646*	0.655*	0.763‡	0.883*	0.993*	0.918*	0.823*	0.851†	0.972	0.302
		-0.026*	-0.017*	0.09*	0.211*	0.073*	-0.001	-0.094*	-0.068*	(0.087)	(0.301)
Italy	β(RS,I,E),σ(I,E)	0.602*	0.581*	0.774*	1.008*	0.685*	0.777*	1.006*	1.087*	0.995	0.998
		-0.046*	-0.133*	0.043*	0.326*	0.003	-0.114*	0.062*	0.298*	(0.000)	(0.000)
Netherlands	β(RS,I,E),σ(I,E)	0.756*	0.867*	0.762*	1.033*	0.741*	0.706*	0.864*	0.927*	0.980	0.984
		-0.005	0.044*	-0.031*	0.184*	-0.065*	-0.114*	0.051*	0.103*	(0.000)	(0.000)
Norway	β(RS,I,E),σ(I,E)		0.816*	0.736*	0.905*		0.904*	0.824*	0.97*	0.985	0.985
			0.135*	-0.1*	0.155*		0.136*	-0.099*	0.133*	(0.000)	(0.000)
Spain	β(RS,I),σ(E)		0.895*	0.923*	1.035*		0.712*	0.998	1.115*	0.991	0.991
			-0.013*	-0.109*	0.011		-0.015*	-0.010	0.153*	(0.312)	(0.391)
Sweden	β(RS,I,E),σ(I,E)		0.759*	1.088*	1.4*		0.665*	0.87*	1.053*	0.986	0.588
		0.006*	-0.166*	0.124*	0.432*	0.027*	-0.119*	0.022*	0.225*	(0.000)	(0.367)
Switzerland	β(RS,I),σ(I,E)	0.576*	0.733*	0.718*	0.87*	0.545*	0.654*	0.753	0.987*	0.994	0.993
		-0.095*	0.013*	-0.088*	0.104*	-0.093*	-0.03*	-0.011	0.261*	(0.000)	(0.000)
UK	β(RS,I,E),σ(I,E)	0.947*	0.851*	0.912*	0.921*	1.432*	1.211*	1.115*	0.858*	0.994	0.942
		-0.016*	0.007*	0.09*	0.099*	0.009*	0.015*	-0.04*	-0.298*	(0.000)	(0.000)
Australia	β(RS,I),σ(I)	0.610*	0.647*	0.664*	0.824*	1.068*	1.09*	0.942*	0.929*	0.923	0.967
		-0.057*	-0.033*	-0.006	0.167*	0.000	0.016*	0.003	-0.083*	(0.000)	(0.000)
Hong Kong	β(RS,I,E),σ(I,E)	0.800*	0.782*	0.879*	0.956*	0.863*	0.969*	1.208*	1.146	0.962	0.977
		0.001	-0.013*	0.026*	0.152*	-0.199*	-0.083*	0.009	0.074*	(0.000)	(0.000)
New Zealand	β(RS,I,E),σ(I,E)		0.591*	0.584*	0.473*		0.654*	0.47*	0.677*	0.200	0.456
			-0.001	-0.014*	-0.125*		0.046*	0.057*	0.265*	(0.303)	(0.398)
Singapore	β(RS,I,E),σ(I,E)	0.729	0.662*	0.622*	0.776*	0.534*	0.546*	0.559*	0.802*	0.944	0.971
		0.093*	0.041*	-0.036*	0.125*	-0.025*	0.011*	-0.034*	0.219*	(0.000)	(0.000)
Canada	β(RS,I,E),σ(I,E)	0.674*	0.753*	0.729*	0.988*	0.442†	0.523*	0.411*	-0.026*	0.980	0.966
		-0.055*	-0.02*	0.04*	0.287*	-0.019*	0.061*	-0.049*	-0.487*	(0.000)	(0.000)

Table 4: Volatility Specification and Significance Tests

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This table reports the results for the volatility specification of the selected model. Panel A shows the results for the regions, Panel B for the industries, and Panel C for the countries. The column ‘Model’ shows whether the specifications for the market betas (‘β’) and conditional variances (‘σ’) contain a RS component and/or Instruments (’I’) and/or a Euro dummy (’E’). The columns ‘Trade’, ‘Align’ and ‘Euro’ show the coefficients of respectively the two structural instruments and the Euro dummy in the volatility specification (if applicable). The next four columns show the coefficients for the Asymmetric GARCH specification. The last part of the table shows the total implied volatility (first line) and its structural component (second line, italics) over different subperiods. *, †, and ‡ indicate significance at respectively the 1, 5, and 10 percent level for the t-test whether subperiod volatilities are equal to the full-period volatilities (first line) and the t-test whether the structural component is equal to zero (second line). The tests are corrected for autocorrelation. The volatility numbers are annualized.

Panel A

Region	**Model**	**Trade**	**Align**	**Euro**	**Const**	**ARCH**	**GARCH**	**Asym**	**73-79**	**80-89**	**90-99**	**00-07**
World	State 1	-	-	-	0.007	0.571	0.008	0.010	0.128*	0.128*	0.124*	0.142*
					(0.001)	(0.000)	(0.383)	(0.391)				
	State 2	-	-	-	0.017	0.744	-0.194	0.271				
					(0.000)	(0.000)	(0.000)	(0.000)				
Europe	β(RS,I),σ(I,E)	-0.155	0.357	-0.475	0.003	0.845	0.059	-0.016	0.112*	0.105*	0.086*	0.079*
		(0.115)	(0.000)	(0.072)	(0.000)	(0.000)	(0.000)	(0.308)	*0.016**	*0.009**	*-0.014**	*-0.018**
Pacific	β(RS,I),σ(I,E)	-0.445	0.075	-0.880	0.004	0.866	0.068	0.024	0.192*	0.18*	0.143*	0.105*
		(0.067)	(0.369)	(0.066)	(0.000)	(0.000)	(0.000)	(0.269)	*0.018**	*0.008**	*-0.014**	*-0.026**
Japan	β(RS,I),σ(I,E)	-0.471	1.194	2.168	0.006	0.763	0.031	0.096	0.149	0.123*	0.147	0.17*
		(0.107)	(0.001)	(0.024)	(0.000)	(0.000)	(0.050)	(0.027)	*0.016**	*-0.02**	*0.013**	*0.027**
US	β(RS,I),σ(I,E)	0.019	0.855	-0.414	0.004	0.781	0.038	0.048	0.062*	0.087*	0.087*	0.06*
		(0.385)	(0.000)	(0.009)	(0.000)	(0.000)	(0.051)	(0.114)	*-0.017**	*0.000*	*0.009**	*-0.017**

Panel B

Industry	Model	Trade	Align	Euro	Const	ARCH	GARCH	Asym	73-79	80-89	90-99	00-07
OILGS	β(RS,I),σ()				-0.001	0.929	0.054	-0.004	0.108*	0.134*	0.114*	0.146*
					(0.294)	(0.000)	(0.000)	(0.389)				
CHMCL	β(RS),σ()				0.002	0.848	0.094	-0.003	0.074*	0.07*	0.071*	0.094*
					(0.004)	(0.000)	(0.000)	(0.396)				
BRESR	β(RS,E),σ(I,E)			2.083	0.004	0.773	0.092	0.027	0.099*	0.1*	0.107*	0.137*
				(0.009)	(0.000)	(0.000)	(0.000)	(0.272)				0.033*
CNSTM	β(RS,I),σ(I)	-0.369	0.421		0.004	0.743	0.122	-0.032	0.114*	0.099	0.09*	0.091*
		(0.048)	(0.012)		(0.000)	(0.000)	(0.000)	(0.257)	0.021*	-0.004*	-0.009*	-0.016*
INDGS	β(RS,I),σ(I,E)	-0.078	0.082	0.118	-0.001	0.875	0.004	0.057	0.049†	0.049*	0.04*	0.048*
		(0.001)	(0.004)	(0.026)	(0.000)	(0.000)	(0.384)	(0.008)	0.002*	0.008*	-0.003*	-0.001
AUTMB	β(RS,I),σ()				0.003	0.871	0.057	0.000	0.107‡	0.103†	0.097*	0.115*
					(0.015)	(0.000)	(0.001)	(0.399)				
FDBEV	β(RS,I,E),σ(E)				-0.001	0.897	0.063	0.029	0.056*	0.06*	0.068*	0.097*
					(0.112)	(0.000)	(0.000)	(0.243)				
PERHH	β(RS,I,E),σ(I,E)	-0.119	0.129		0.002	0.876	0.039	0.025	0.073‡	0.072	0.064*	0.078*
		(0.081)	(0.036)		(0.000)	(0.000)	(0.027)	(0.279)	0.009*	0.000	-0.004*	-0.007*
HLTHC	β(RS,I,E),σ(I,E)	-0.008	0.073		-0.001	0.899	0.026	0.061	0.058*	0.058*	0.067†	0.087*
		(0.395)	(0.020)		(0.000)	(0.000)	(0.033)	(0.021)	-0.005*	-0.001*	0.005*	0.007*
RTAIL	β(RS,I),σ()				0.002	0.898	0.018	0.068	0.110†	0.101*	0.093*	0.121*
					(0.000)	(0.000)	(0.196)	(0.010)				
MEDIA	β(RS,I),σ(I)		0.161		0.002	0.901	0.052	0.008	0.118*	0.086*	0.072*	0.087†
			(0.001)		(0.000)	(0.000)	(0.000)	(0.379)	0.008*	0.006*	-0.004*	-0.008*
TRLES	β(RS,I),σ(I,E)			-0.600	0.004	0.849	0.035	0.043	0.127*	0.106*	0.100*	0.108
				(0.004)	(0.000)	(0.000)	(0.031)	(0.087)	0.000	0.003*	-0.004*	-0.018*
TELCM	β(RS,I,E),σ(I,E)				0.002	0.873	0.057	0.018	0.105*	0.125*	0.093*	0.125*
					(0.018)	(0.000)	(0.000)	(0.279)				
UTILS	β(RS,I),σ(I,E)				-0.002	0.873	0.109	-0.042	0.091	0.104*	0.076*	0.097*
					(0.006)	(0.000)	(0.000)	(0.122)				
BANKS	β(RS,I),σ(I)		0.131		-0.002	0.883	0.070	0.010	0.097‡	0.115*	0.082*	0.075*
			(0.012)		(0.001)	(0.000)	(0.000)	(0.374)	0.006*	0.003*	-0.003*	-0.011*
INSUR	β(RS),σ()				-0.002	0.900	0.056	0.004	0.076*	0.077*	0.075*	0.092*
					(0.000)	(0.000)	(0.000)	(0.392)				
FINSV	β(RS,I),σ(I)		0.146		-0.002	0.899	0.076	-0.029	0.100	0.124*	0.093*	0.068*
			(0.012)		(0.002)	(0.000)	(0.000)	(0.206)	0.011*	0.009*	-0.003*	-0.006*
TECNO	β(RS,I,E),σ(E)				0.001	0.939	0.037	0.022	0.119*	0.123*	0.136	0.161*
					(0.308)	(0.000)	(0.000)	(0.155)				

Panel C

Country	Model	Trade	Align	Euro	Const	ARCH	GARCH	Asym	73-79	80-89	90-99	00-07
Austria	β(RS,I,E),σ(E)			-1.801	0.004	0.825	0.101	0.029	0.123*	0.181*	0.154	0.128*
				(0.000)	(0.000)	(0.000)	(0.001)	(0.269)				*-0.031**
Belgium	β(RS,I,E),σ(I,E)	-0.085	-0.003	-0.194	0.002	0.932	0.027	0.020	0.125	0.151*	0.112*	0.104*
		(0.321)	(0.398)	(0.006)	(0.000)	(0.000)	(0.015)	(0.195)	*0.007**	*0.004**	*-0.004**	*-0.013**
Denmark	β(RS,I),σ(I,E)	-0.142	0.455	-0.405	0.003	0.915	0.052	-0.009	0.204*	0.177*	0.127*	0.127*
		(0.374)	(0.041)	(0.007)	(0.000)	(0.000)	(0.007)	(0.377)	*0.018**	*0.001*	*-0.015**	*-0.017**
Finland	β(RS,I,E),σ(I)	2.573	1.567		0.004	0.945	-0.030	0.043		0.17*	0.232	0.234
		(0.188)	(0.000)		(0.188)	(0.000)	(0.349)	(0.131)		*0.013**	*-0.025**	*-0.016**
France	β(RS,I,E),σ(I,E)	-0.232	0.211	-0.350	-0.003	0.873	0.070	0.026	0.171*	0.17*	0.099*	0.06*
		(0.311)	(0.000)	(0.000)	(0.000)	(0.000)	(0.003)	(0.263)	*0.014**	*0.008**	*-0.009**	*-0.02**
Germany	β(RS,I,E),σ(I,E)	-0.284	0.265	-0.017	-0.003	0.840	0.050	0.038	0.118*	0.123*	0.091*	0.08*
		(0.212)	(0.000)	(0.000)	(0.000)	(0.000)	(0.016)	(0.166)	*0.013**	*0.011**	*-0.003**	*-0.017**
Ireland	β(RS,I,E),σ(I,E)	-0.472	0.853	0.324	0.004	0.895	0.043	0.015	0.211*	0.181*	0.131*	0.138*
		(0.020)	(0.003)	(0.054)	(0.000)	(0.000)	(0.025)	(0.335)	*0.024**	*0.011**	*-0.02**	*-0.01**
Italy	β(RS,I,E),σ(I,E)	-0.200	0.858	-2.069	-0.006	0.873	0.089	-0.026	0.226*	0.232*	0.185*	0.081*
		(0.359)	(0.154)	(0.013)	(0.000)	(0.000)	(0.004)	(0.232)	*0.019**	*0.016**	*-0.017**	*-0.04**
Netherlands	β(RS,I,E),σ(I,E)	-0.140	0.194	-0.104	0.002	0.868	0.064	-0.001	0.109*	0.107*	0.082*	0.067*
		(0.215)	(0.029)	(0.000)	(0.000)	(0.000)	(0.001)	(0.399)	*0.009**	*0.011**	*-0.004**	*-0.015**
Norway	β(RS,I,E),σ(I,E)	0.351	1.773	-0.809	0.005	0.884	0.021	0.058		0.251*	0.189*	0.166*
		(0.234)	(0.000)	(0.000)	(0.000)	(0.000)	(0.166)	(0.025)		*0.018**	*-0.02**	*-0.022**
Spain	β(RS,I),σ()				-0.002	0.978	-0.029	0.057		0.138*	0.132*	0.089*
					(0.231)	(0.000)	(0.396)	(0.397)				
Sweden	β(RS,I,E),σ(I,E)	-0.820	-0.037	-0.041	-0.004	0.871	0.020	0.024		0.195*	0.182*	0.137*
		(0.097)	(0.397)	(0.070)	(0.000)	(0.000)	(0.253)	(0.270)	*-0.014**	*0.004**	*0.009**	*-0.022**
Switzerland	β(RS,I),σ(I,E)	-0.146	0.166	-0.294	0.003	0.900	0.054	-0.028	0.132*	0.11*	0.105*	0.081*
		(0.270)	(0.018)	(0.000)	(0.000)	(0.000)	(0.000)	(0.164)	*0.011**	*-0.004**	*0.001**	*-0.016**
UK	β(RS,I,E),σ(I,E)	0.039	0.321	-0.734	0.004	0.780	0.016	0.111	0.129*	0.091	0.069*	0.063*
		(0.371)	(0.021)	(0.000)	(0.000)	(0.000)	(0.282)	(0.000)	*0.012**	*0.004**	*-0.015**	*-0.017**
Australia	β(RS,I),σ(I)	-0.428	0.358		0.006	0.664	0.040	0.043	0.124*	0.12*	0.107*	0.087*
		(0.109)	(0.095)		(0.000)	(0.000)	(0.093)	(0.190)	*0.012**	*0.005**	*-0.009**	*-0.023**
Hong Kong	β(RS,I,E),σ(I)	0.248	0.428		-0.003	0.816	0.081	0.061	0.272*	0.235*	0.125*	0.087*
		(0.222)	(0.000)		(0.015)	(0.000)	(0.001)	(0.085)	*0.019**	*0.002**	*-0.011*† *	*-0.023*† *
New Zealand	β(RS,I,E),σ(I)	-3.665	6.654		-0.007	0.659	0.026	0.020		0.197*	0.159*	0.129*
		(0.067)	(0.072)		(0.003)	(0.000)	(0.330)	(0.379)		*0.013**	*0.038**	*0.007**
Singapore	β(RS,I,E),σ(I,E)	-0.657	-0.079	-0.825	-0.004	0.838	0.052	0.070	0.233*	0.198*	0.151*	0.117*
		(0.078)	(0.380)	(0.002)	(0.000)	(0.000)	(0.018)	(0.024)	*0.020**	*0.011**	*-0.018**	*-0.026**
Canada	β(RS,I,E),σ(I,E)	-0.051	0.480	-0.413	0.004	0.807	0.084	-0.013	0.124*	0.113*	0.103*	0.106*
		(0.388)	(0.027)	(0.000)	(0.000)	(0.000)	(0.000)	(0.360)	*0.015**	*-0.009**	*-0.007**	*-0.014**

Table 5: Average Residual Correlations and Market-Weighted Average Betas

Panel A shows average correlations between and across region, country, and industry-specific shocks over the full sample period and over different subsamples (for the latter). It further reports correlations of all asset-specific shocks with world market shocks. To the extent that the chosen structural regime-switching volatility spillover model is capable of modeling cross-asset correlations, we expect residual correlations both within and between assets to be close to zero. Panel B displays the value-weighted average of respectively regional betas with respect to the global market(β^{w}_{reg}), country betas with respect to the global market (β^{w}_{c}), country betas with respect to the regional market (β^{reg}_{c}), and industry betas with respect to the global market (β^{w}_{gi}), over the full sample period and over different subperiods. For the full sample period, the cross-sectional variation in the betas is also reported (*Stdev*).

A. Average Residual Correlations								
	World	**Region**	**Country**	**Industry**	**73-79**	**80-89**	**90-99**	**00-07**
World	1.00	0.02	0.00	0.00	-	-	-	-
Region	0.02	-0.17	0.03	0.01	-0.06	-0.09	0.04	-0.09
Country	0.00	0.01	0.01	0.00	0.18	0.02	0.08	0.04
Industry	0.00	0.03	0.00	0.01	-0.01	0.04	0.03	0.07
B. Market-Weighted Average Betas								
	Mean		**Stdev**		**73-79**	**80-89**	**90-99**	**00-07**
β^{w}_{reg}	0.983		0.044		0.981	0.985	0.987	0.975
β^{w}_{c}	0.978		0.051		0.992	0.970	0.977	0.975
β^{reg}_{c}	0.979		0.015		0.986	0.982	0.982	0.965
β^{w}_{gi}	0.982		0.024		0.989	0.985	0.982	0.973

Figure 1: Smoothed Probabilities of the World Volatility Regimes

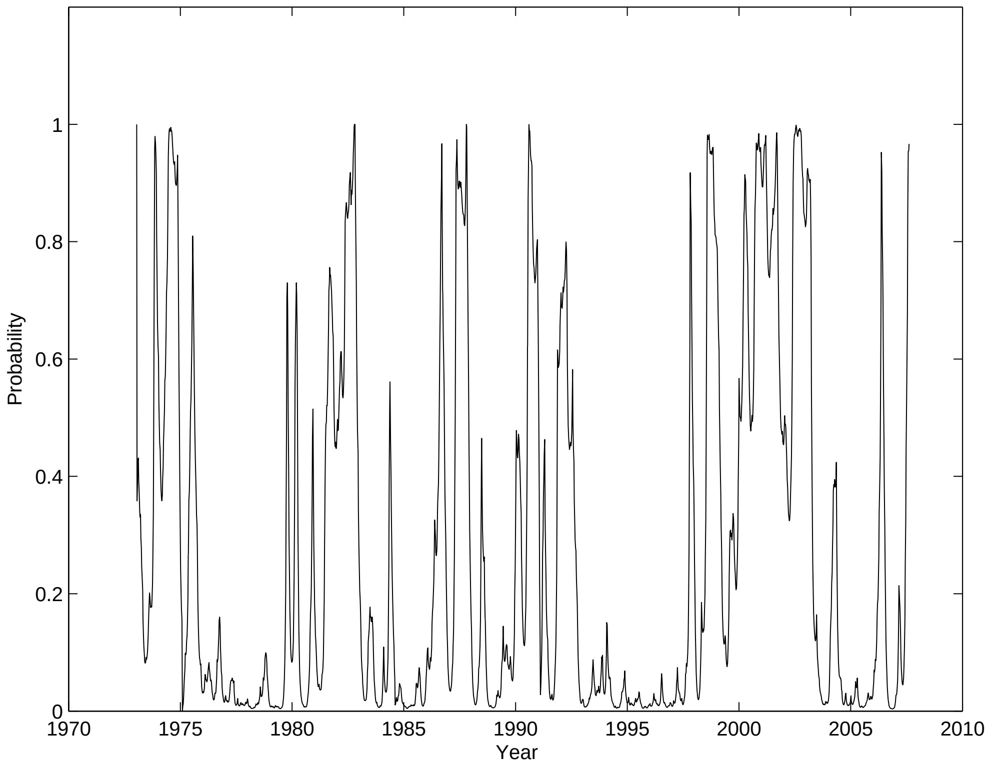
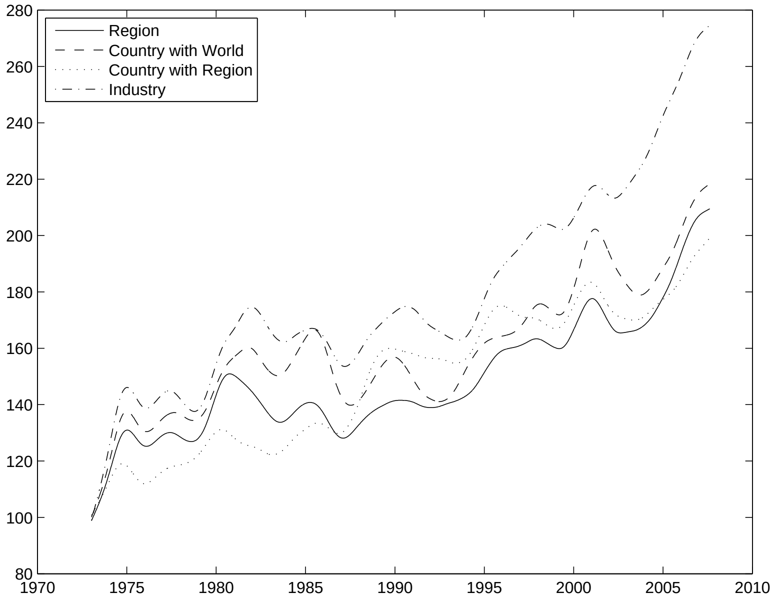


Figure 1 plots the smoothed probabilities that the world equity returns are in the high volatility state. The probabilities are obtained from estimating the regime-switching Asymmetric GARCH model outlined in II.A.

Figure 2: Evolution of Structural Instruments over Time for Regions, Countries and Industries

Panel A: Trade Integration



Panel B: Misalignment

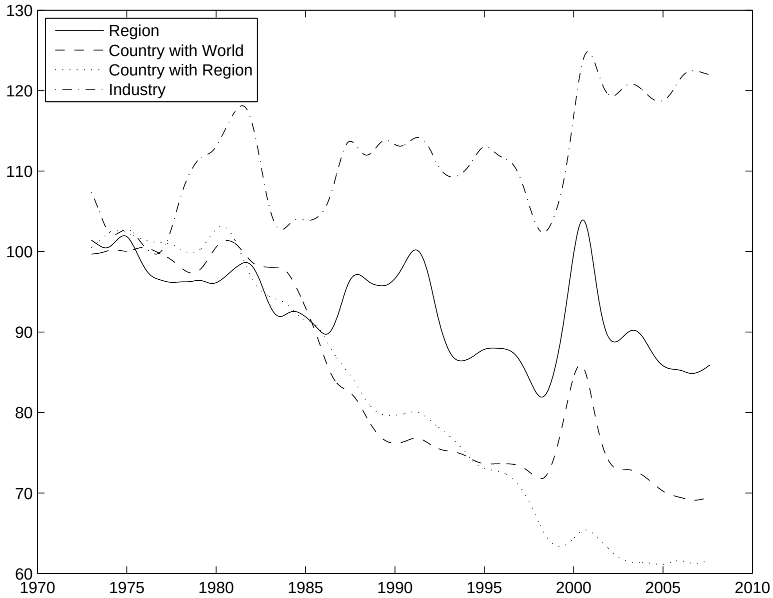
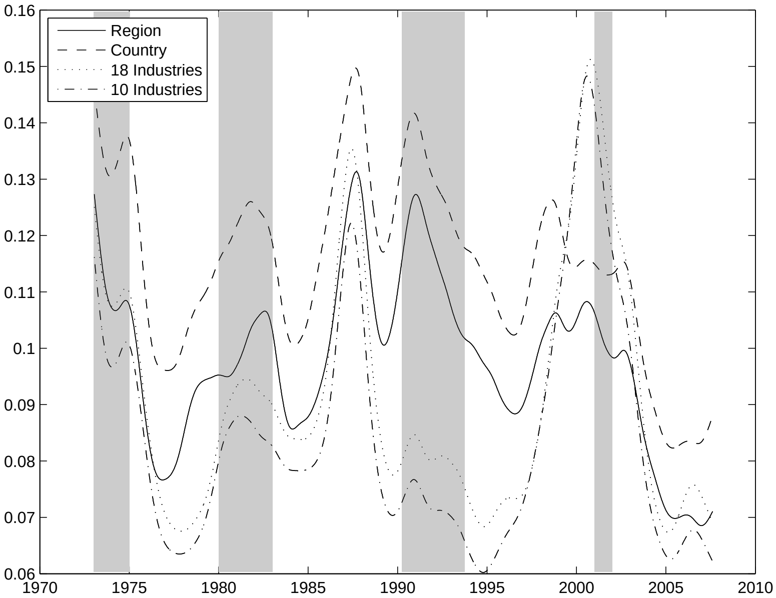


Figure 2 plots a backward 52-week moving average of trade integration (Panel A) and misalignment (Panel B) at the regional, country, and industry level. At the country level, we differentiate between trade integration and misalignment with respect to the world (excluding the region) and with respect to the region. The measures are constructed as indices starting at 100 at the beginning of the sample and are averaged over the specific levels.

Figure 3: Idiosyncratic Volatility Aggregated over Regions, Countries and Industries

Panel A: Including TMT



Panel B: Excluding TMT

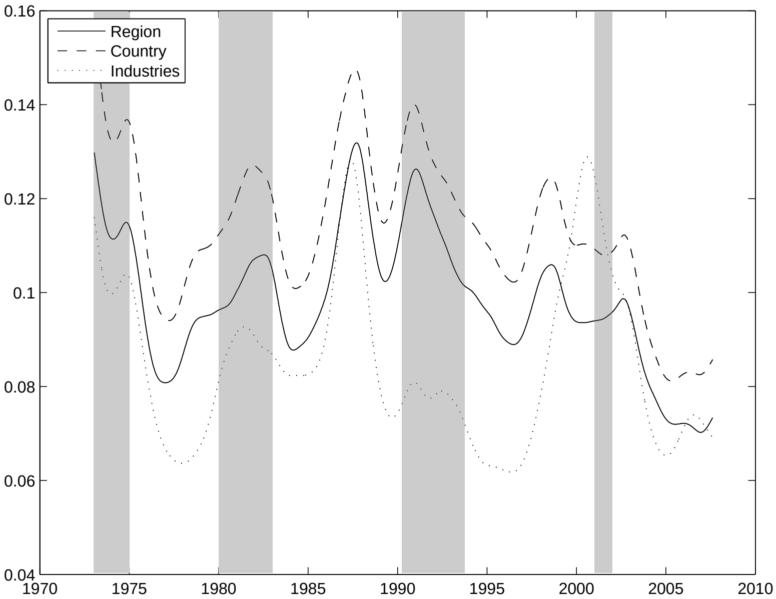


Figure 3 plots a backward 52-week moving average of the average idiosyncratic volatilities at the regional, country, and industry level, both including (Panel A) and excluding the TMT industries (Panel B). The latter are respectively the Telecom, Media, and Information Technology (both Software and Hardware) industries. We use market capitalization weights to average over the asset-specific volatilities. The latter are obtained by estimating the structural regime-switching model developed in Section II.A. World recessions are shaded in gray to illustrate cyclical movements in volatility. The recessions are identified as the periods from the peaks to the throughs of the detrended world GDP.

Figure 4: Volatility Decomposition for Aggregated European Countries

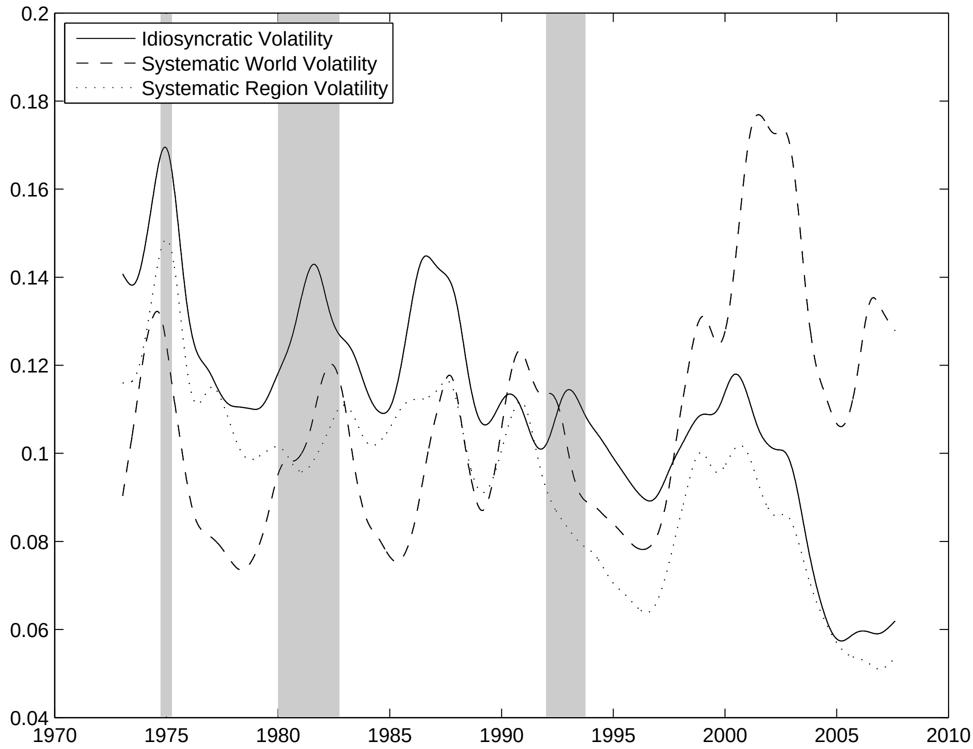
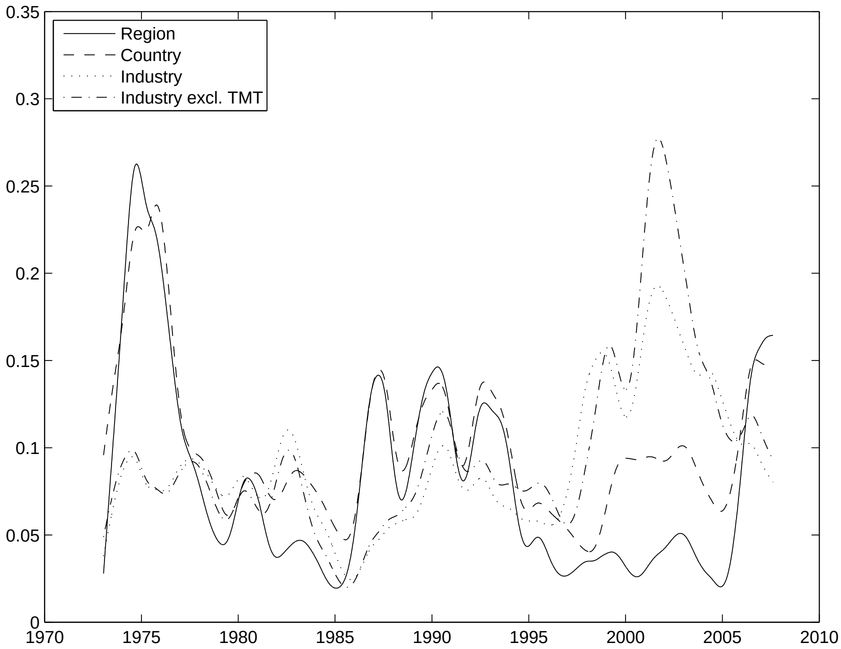


Figure 4 plots a backward 52-week moving average of the average idiosyncratic volatilities across European countries. The figure also plots the contibution of global and region-specific volatility to the total country volatility (in absolute value). We use market capitalization weights to average over the asset-specific volatilities. All idiosyncratic volatilities are obtained by estimating the structural regime-switching model developed in Section II.A. CEPR-dated recessions for Europe are shaded in gray to illustrate cyclical movements in volatility.

Figure 5: Quantification of Bias in Measures of Average Idiosyncratic Volatility

Panel A: Bias due to Assumption of Unit Betas



Panel B: Bias due to Assumption of Constant Betas

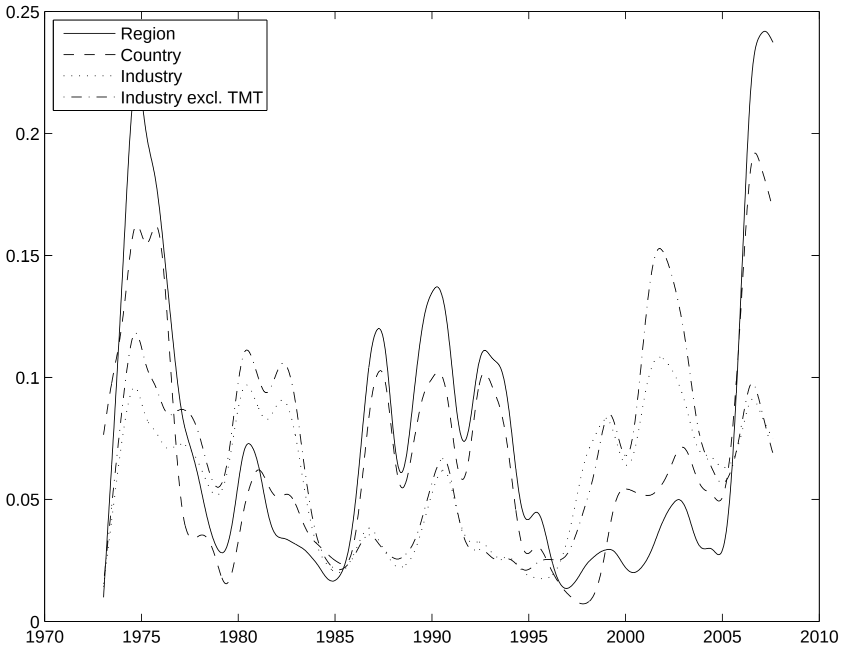
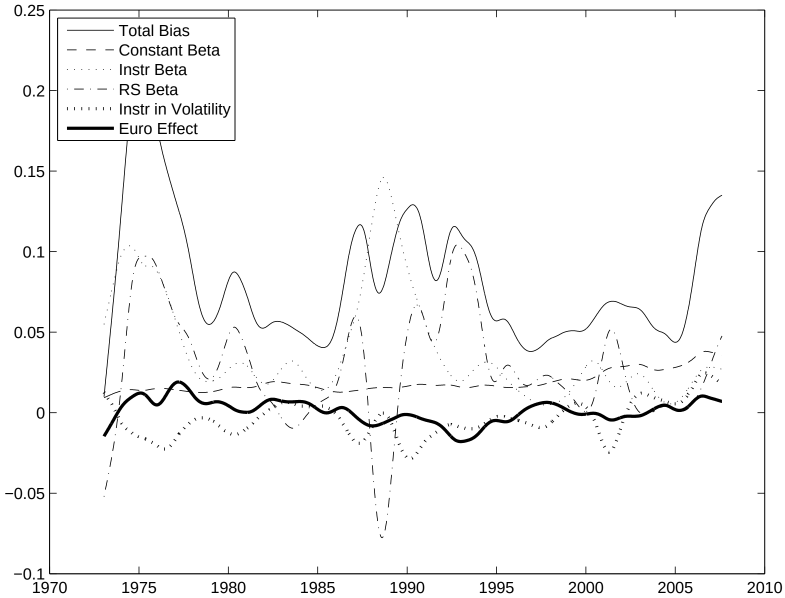


Figure 5 plots the bias in the measures of average idiosyncratic risk resulting from imposing unit betas (Panel A) and constant betas (Panel B) instead of time-varying betas. We refer to Section V.B. and to an appendix available from the authors' websites for the exact specification and derivation of the biases.

Figure 6: The Decomposition of the Bias in the Nonsystematic Risk Component

Panel A: Decomposition of Bias at Country Level



Panel B: Decomposition of Bias at Industry Level

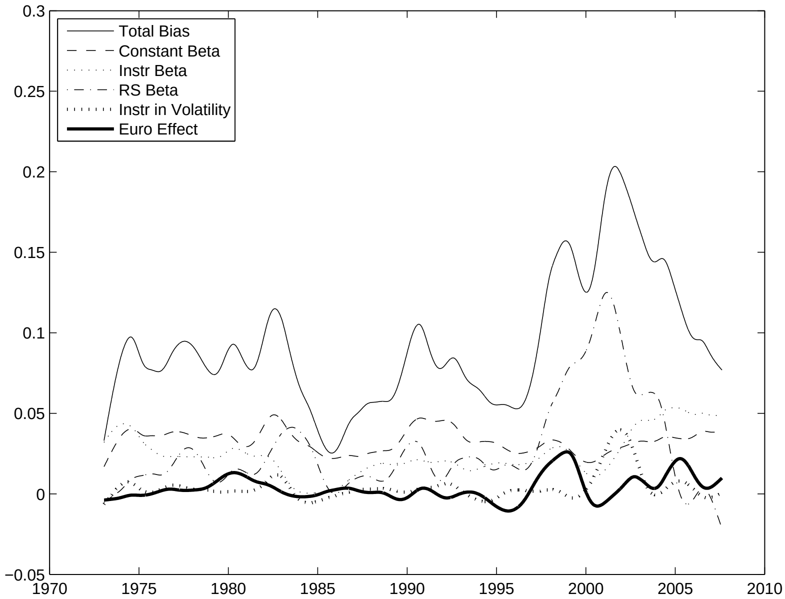


Figure 6 shows the individual contribution of adding various complexities to the volatility spillover model to a reduction in the total bias caused by imposing unit global market exposures. We respectively show the contribution of allowing betas to be constant instead of being unity, of allowing structural instruments instruments in the betas (relative to the constant beta case), of allowing regime-switches in the betas (relative to the beta specification with instruments), and finally of also allowing for structural shifts in the asset-specific volatility specification (relative to model with time-varying betas and an AGARCH volatility specification). The individual contributions sum up to the toal bias. We report the decomposition at the country level (Panel A) and at the industry level (Panel B).

Figure 7: Value-Weighted Model-Implied Correlations over Regions, Countries and Industries

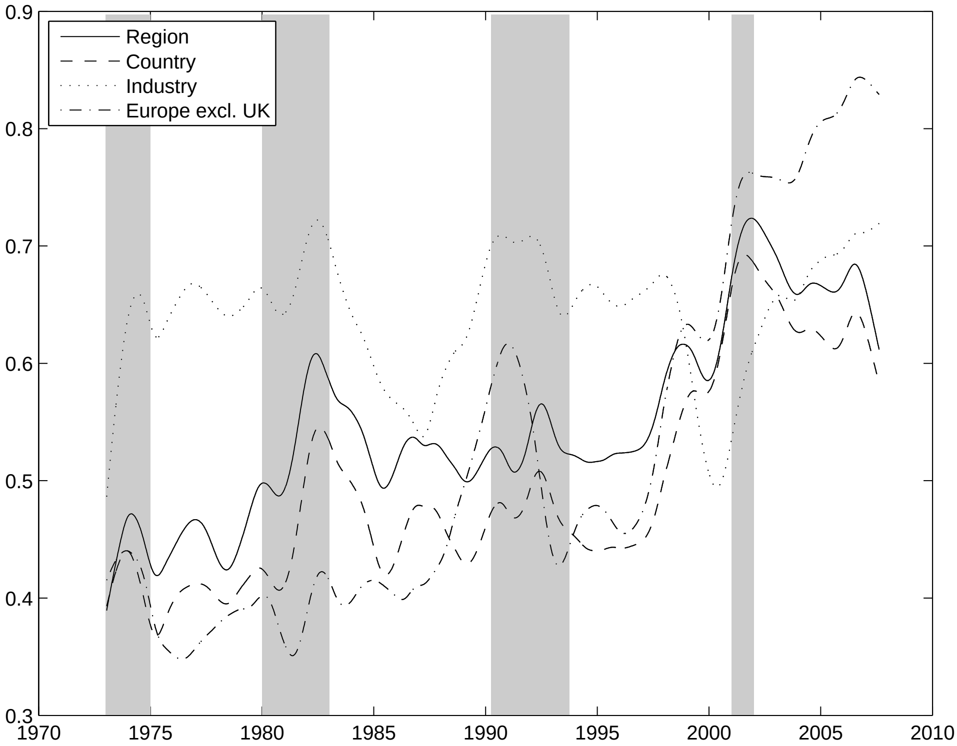


Figure 7 reports the average model-implied cross-regional, cross-industry, and cross-country correlations over time. For the cross-country correlations, we distinguish between all countries and the European countries. World recessions are shaded in gray to illustrate cyclical movements in volatility. The recessions are identified as the periods from the peaks to the throughs of the detrended world GDP.

Figure 8: Quantification of Bias in Model-Implied Correlations over Regions, Countries and Industries

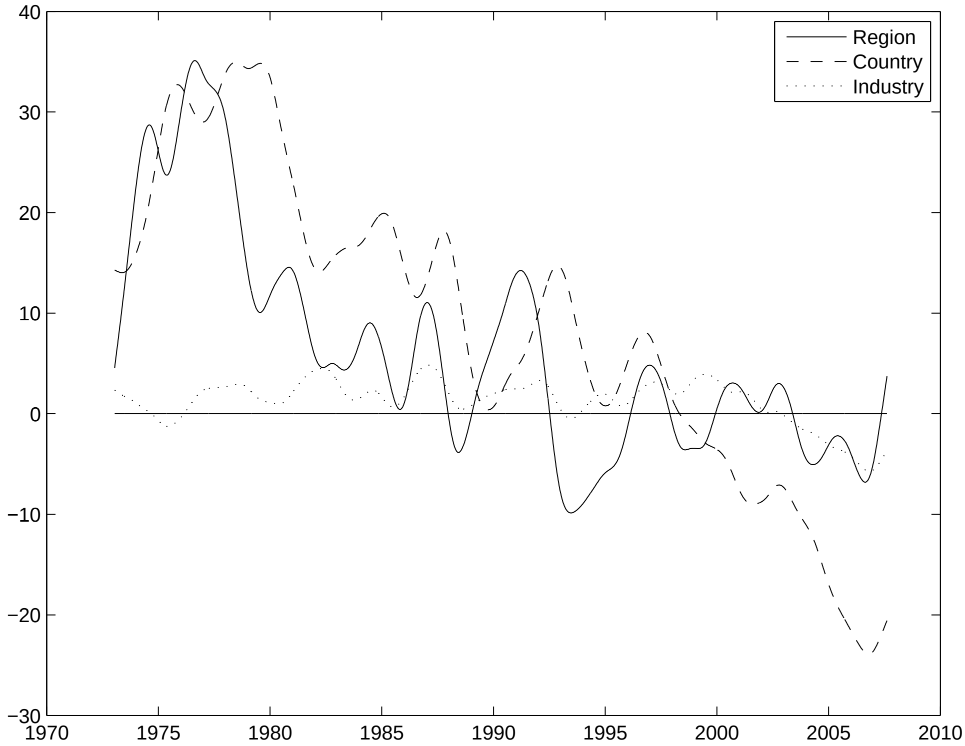


Figure 8 plots the bias in the equally-weighted model-implied correlations over regions, countries and industries from imposing constant betas instead of time-varying betas and not allowing for structural changes in the idiosyncratic volatility.